Government Policy Uncertainty and the Yield Curve*

Markus Leippold[†]   Felix H. A. Matthys[‡]

August 23, 2017

Abstract

We study the impact of government policy uncertainty on the term structure of nominal interest rates. We develop a general equilibrium model, in which both the real as well as the nominal side of the economy are driven by government uncertainty shocks. Our affine yield curve model captures both the shape of the interest rate term structure as well as the hump-shape of bond yield volatilities. Our theoretical predictions are strongly supported by the data. Higher government policy uncertainty leads to a significant decline in yield levels, induces a hump-shaped increase in bond yield volatility, and increases bond risk premia, especially for longer maturities.

JEL classification: G01, G12, G14, G18
Key Words: Term structure modeling, yield volatility curve, policy uncertainty, bond risk premia

---

*This paper benefited greatly from discussions with Yacine Aït-Sahalia, Caio Almeida, Markus Brunnermeier, Olivier Darmouni, Jérôme Detemple, Itamar Drechsler, Darell Duffie, Valentin Haddad, Oleg Itskhoki, Jakub Jurek, Andrew Karolyi, Jean-Charles Rochet, Christopher Sims, David Srear, Adi Sunderam, Josef Teichmann, Fabio Trojani, and Wei Xiong. Special thanks go to our discussants Philippe Mueller, Anna Cieslak, and Michael Weber. For helpful comments, we would like to thank the seminar participants of the 2015 SAFE Asset Pricing Workshop in Frankfurt, the Finance and Math Seminar ETH and University of Zurich, the 12th Doctoral Workshop in Finance at Gerzensee, the Princeton Student Research Workshop, the Financial Mathematics Seminar at ORFE Princeton University, the Recent Advances in fixed-income research and implications for monetary policy conference (Federal Reserve Board San Francisco and Bank of Canada), AEA Meetings San Francisco, EEA Meetings Geneva, Bacelona GSE Summer Forum, Bank of England, Central Bank of Mexico, and ITAM. Financial support from the Swiss Finance Institute (SFI), Bank Vontobel, the Swiss National Science Foundation and the National Center of Competence in Research "Financial Valuation and Risk Management" is gratefully acknowledged.

†University of Zurich, Department of Banking and Finance, Plattenstrasse 14, 8032 Zurich, Switzerland; markus.leippold@bf.uzh.ch.

‡Instituto Tecnológico Autónomo de México (ITAM), Dept. of Business Administration E-mail: felix.matthys@itam.mx.

1 Introduction

Political uncertainty is a ubiquitous byproduct of policymaking. Given the far-reaching impact of government decisions on the economy, government policy uncertainty has important consequences for long-run economic growth, asset prices, and welfare. In uncertain times, investors commonly seek shelter in safe haven assets such as treasury bonds. Consequently, government policy uncertainty has a direct impact on the term structure of interest rates. For this reason, we examine the link between government policy uncertainty and yields, the volatility curve, and bond risk premia. We develop a general equilibrium model, in which policy uncertainty shocks affect both the real and the nominal side of the economy.

Our model economy consists of a real and a monetary sector with an infinitely lived representative agent. While the central bank controls money supply on the basis of a Taylor rule, the government influences the real sector of the economy using a set of different policy instruments.[1] To introduce government policy uncertainty, we let the representative agent form his own expectation about what policy the government will select. The discrepancy between his expectation and the actual implemented policy constitutes our measure of government policy uncertainty.[2] Therefore, if the government choses a widely anticipated policy, its impact on the economy is marginal. Contrarily, if the selected policy deviates considerably from general expectations, it not only triggers large reactions in financial markets but also adversely affects the economy.

We find that government policy uncertainty impacts both the real and the nominal side in various different ways. For the real side, it not only negatively affects the long-run growth path of production, but it also increases its volatility and therefore leads to a worsening of economic growth prospects, which are fundamental to the agent's consumption and investment allocation decisions.[3]

[1] To keep our analysis focused, we use a reduced form approach when modeling government policies.

[2] In Croce, Kung, Nguyen, and Schmid (2012) and Croce, Nguyen, and Schmid (2012), (tax) policy uncertainty is defined as the difference between the actual and expected tax rate. Here, we take the squared difference as a measure of uncertainty. Following the terminology used in Scotti (2016), the squared difference may serve as a measure of 'policy uncertainty,' the focus of our paper, while the simple difference (positive or negative) may be interpreted as 'policy surprise.'

[3] Our model has some similarities to the long-run risk model of Bansal and Yaron (2004). However, the key distinctive difference is that the long-run growth component and the market price of output risk are both driven by the same underlying risk factor, namely policy uncertainty. This modeling assumption is supported by empirical results by Baker and Bloom (2013), who show that both first and second order uncertainty shocks explain a significant proportion of GDP growth. Furthermore, our setting can also be compared to the literature on real business cycle analysis. For

This observation is in line with Bloom (2009), who argues that productivity growth falls, because higher uncertainty causes firms to temporarily pause their investments. Hence, an increase in policy uncertainty renders capital investments more risky, which will eventually induce investors to favor safe assets such as government bonds. Such a flight-to-quality behavior will raise government bond prices and therefore drives down their yields.

For the nominal side of the economy, we assume that the central bank conducts its monetary policy through a Taylor rule, which is affected by policy uncertainty through two different channels. The first one is a direct effect, where an increase in uncertainty causes the central bank to increase its money supply, driving down interest rates in the economy. The second channel is indirect. As policy uncertainty drives both capital accumulation and the price level in the economy, which are the two main variables entering the Taylor rule, it indirectly influences the central bank's money provision.

Our general equilibrium model allows us to study the numerous ways how policy uncertainty affects not only the level of interest rates, but also how the level and shape of the term structure of bond yield volatilities react in response to an uncertainty shock. Although our model belongs to the class of affine models as introduced by Duffie and Kan (1996), we can replicate the typical hump-shape of the volatility term structure, caused by the empirical observation that volatility tends to be highest around the two-year maturity bucket.[4] The key mechanism leading to this result is that policy uncertainty negatively affects the long-run growth path of productivity, which translates into a hump-shaped curve in the term structure of volatility. With this amplification mechanism we can also explain the 'excess bond yield volatility puzzle' that empirical bond yields cannot be reproduced by standard affine models of the term structure of interest rates (see Shiller (1979) and Piazzesi and Schneider (2006)).[5] Finally, we allow government policy shocks to carry a risk premium.

instance, shocks to trend growth exhibit fundamentally different effects on the (real) economy as opposed to transitory fluctuations. The agents or country's reaction to temporary shocks is to borrow in the short run to smooth out consumption. However, if the shock is more persistent, the long-run consumption level has to be adjusted as borrowing for an infinite time horizon is not feasible.

[4]Matching the hump-shape in the volatility data using affine term structure models is a challenging task. For example, as Buraschi and Jiltsov (2005) argue, even with a more flexible specification of market price of risk, their model is not able to replicate the hump in the term structure of bond volatility. Furthermore, Malkhozov, Mueller, Vedolin, and Venter (2016) introduce a feedback mechanism in which negative convexity increases the market price of risk to match the empirical shape of bond yield volatility. Even though their model is able to replicate the hump, it fails to match the overall level of the term structure of bond volatility.

[5]A possible solution to this problem is to introduce heterogeneous agents who have different prior beliefs about some

Hence, our model accommodates time-varying risk premia in bond returns, which implies that policy uncertainty is a priced risk factor.

Dealing with policy uncertainty, a fundamental question that arises in this context is: What is an appropriate measure for government and monetary policy uncertainty? In our empirical analysis, we rely on the economic policy uncertainty (EPU) index developed by Baker, Bloom, and Davis (2016) as a proxy for aggregate government policy uncertainty. To motivate our choice, Figure 1 plots the relationship between U.S. treasury bond yields and volatilities together with the EPU index for the period of January 1990 to September 2015. The first two prominent spikes of the EPU index are related to the terrorist attacks on the World Trade Center and the 2nd Gulf War. By the end of 2003 and until the outbreak of the financial crisis in 2008, the US economy entered a steady economic growth phase. The EPU declined in the pre-crisis period and started to peak at the onset of the financial crisis and remains at a high level ever since, exhibiting highly volatile behavior. We attribute this observation to the fact that the EPU index captures political uncertainty, which was especially high during the debt-ceiling crisis of 2011 and lasted until late 2013 where some governmental authorities were even forced to suspend their services temporarily.

[Figure 1 about here.]

Figure 1 also shows that the EPU index exhibits a countercyclical pattern with nominal yields. When the EPU index is high, yields tend to go down. This apparent negative relationship is also confirmed by computing the sample correlation coefficient, which ranges between -0.53 and -0.39 for the one year and ten year yield, respectively. These numbers suggest that, as political risk increases, investors seek safer assets and therefore start to shift from stocks to (government) bonds and thereby lowering yields, which is in line with the predictions in Pastor and Veronesi (2013).[6]

[Figure 2 about here.]

fundamental economic variable, such as for instance inflation as in Xiong and Yan (2010) or to introduce time-varying risk preference as in Buraschi and Jiltsov (2007), which are however analytically less tractable.

[6] Using also the EPU index of Baker, Bloom, and Davis (2016), Pastor and Veronesi (2013) show that political uncertainty raises not only the equity risk premium but also the volatilities and correlations of stock returns.

There is increasing evidence that policy uncertainty leads to direct reactions of the central bank authority (see for instance David and Veronesi (2014)). To motivate a link between policy uncertainty and yields for our model design, we estimate pairwise Vector Auto Regressions (VAR) for the EPU with the effective fed funds rate, which we take as proxy for monetary policy. Figure 2 reports the resulting impulse responses for the time period from January 1990 to September 2015. Panel A reveals that a shock to the EPU index leads to a sustained negative impact on the feds funds rate and hence on future monetary policy. This impact remains highly significant up to a time horizon of more than 20 months. In contrast, from Panel B we observe that the impulse response function is flat at zero. Hence, a shock to the short-term rate has no impact on the EPU index.[7] This finding suggests that policy uncertainty shocks drive monetary policy actions, but not the other way around. Hence, the central bank conducts its monetary policy taking into account uncertainty shocks from the real side whereas the central banks interest rate policy does not seem to affect uncertainty.

We provide a theoretical explanation for the empirical observations above. In our model economy, higher real uncertainty lowers productivity growth, which feeds into the monetary policy through our assumption that the central bank controls the money supply growth following a Taylor rule. Hence, the monetary authority's efforts to stabilize growth (and inflation) causes it to react to real uncertainty by lowering the cost of capital. Moreover, since we assume money neutrality, nominal shocks do not have an impact on the real side of the economy. However, the converse is not true. The equilibrium price level growth is driven by the capital accumulation growth, which implies that the nominal side is also driven by shocks from the real side, namely government policy uncertainty shocks. Through these two transmission channels, inflation and capital accumulation growth targeting, the money supply growth becomes a function of government policy uncertainty. Therefore, by letting the central bank react endogenously to deviations from long-run capital growth and inflation targets, we can establish an important link between the real and nominal side of the economy. This link allows government policy uncertainty to affect nominal quantities, which proves to be essential to simultaneously match the term structure of interest rates and its corresponding volatility curve.

[7] Further analyzing the impulse responses from our VAR model, when we add the three month and ten year TB rate, we find that the EPU index has a similar impact along the entire term structure. However, our study shows that the short end of the term structure response more strongly to policy uncertainty shocks than the long end does. We do not report these graphs here, but they can be obtained on request.

Finally, our affine term structure model is not only able to replicate the empirical shape of both the level and the volatility of the yield curve, but it also generates theoretical predictions that we can confirm in an empirical exercise. In particular, we find that higher government policy uncertainty leads to a significant decline in contemporaneous yield levels, induces a hump-shaped increase in bond yield volatility, and positively predicts bond risk premia, especially for longer maturities.

Our paper belongs to the class of general equilibrium models of the term structure.[8] However, this literature abstracts from modeling government policy uncertainty and remains silent about its potential impact on the yield curve. Especially since the European debt crisis starting in 2010 and the 2011 Congress debate about raising the fiscal debt ceiling in the US, policy uncertainty has attracted considerable interest from academia. For instance, Pastor and Veronesi (2012, 2013) develop a general equilibrium model, in which the profitability of firms is driven by government policy, and discuss the impact of policy risk on stock prices. Additionally, several recent asset pricing papers have linked fiscal or tax policy uncertainty to a rise in long-run risk, which depresses innovation and may lead to welfare losses (see, e.g., Croce, Nguyen, and Schmid (2012) and Croce, Kung, Nguyen, and Schmid (2012)) Besides, numerous empirical papers have shown that uncertainty about political outcomes has a significant effect on asset returns and corporate decisions.[9] There is also a large strand of literature trying to infer political risk from government bond yields such as, e.g., Huang, Wu, Yu, and Zhang (2015) who empirically study the relationship between political risk and government bond yields.[10] Hence, we fill a gap in the literature by providing both a theoretical model on how government policy uncertainty may impact the yield curve and an empirical analysis of the hypotheses derived from our model.

While the literature on government policy impacts is sparse but growing, the fundamental link between monetary policy and the term structure of interest rates and volatilities has been studied

[8]This class of models includes, among many others, Wachter (2006), Piazzesi and Schneider (2006), Buraschi and Jiltsov (2007), Gallmeyer, Hollifield, Palomino, and Zin (2007), Bekaert, Engstrom, and Xing (2009), and Bansal and Shaliastovich (2013).

[9]For instance, early studies include Rodrik (1991) and Pindyck and Solimano (1993). More recently, Durnev (2010) and Julio and Yook (2012) document that firms tend to withhold their investment activity prior to national elections. Gulen and Ion (2012) argue, based on the EPU index of Baker, Bloom, and Davis (2016), that policy uncertainty substantially reduces firm and industry level investment. Boutchkova, Doshi, Durnev, and Molchanov (2012) take the analysis further and show that some industries are more sensitive to political uncertainty than others. Some further related articles analyzing the relationship between political uncertainty and asset returns include Belo, Gala, and Li (2013), Bialkowski, Gottschalk, and Wisnieski (2008) and Bond and Goldstein (2015).

[10]For an overview of this literature, we refer to Bekaert, Harvey, Lundblad, and Siegel (2014).

more extensively.[11] Despite recent attention brought to modeling the impact of policy uncertainty on asset prices, the papers mentioned above either address the empirical link between government bond yields and policy uncertainty or focus on the theoretical impact that a given government policy has on stock returns. Hence, our paper fills a gap in that it provides both a theoretical framework and an empirical analysis on the impact of policy uncertainty on the nominal yield curve and on bond yield volatility.

The reminder of the paper is organized as follows. Section 2 presents the model. Section 3 discusses the impact of government as well as monetary policy uncertainty on the term structure of nominal interest rates, the yield volatility curve and the bond risk premium. Section 4 summarizes our empirical results and Section 5 concludes.

2 The baseline model economy

To introduce uncertainty about the government's policy, we proceed as follows. For a given government policy, $X_{i,t}$, which could be, e.g., the corporate tax rate fluctuation as in Croce, Kung, Nguyen, and Schmid (2012), our representative agent forms his own expectation the government is going to select. The squared deviation from his expectation and the actual policy implemented by the government is what we refer to as government policy uncertainty for a given policy X_i. Assuming that the government can implement n different policies, we can define an aggregate government policy uncertainty (GPU) index as

$$v_t := \sum_{i=1}^{n} \left(X_{i,t} - \mathbb{E}[X_{i,t}] \right)^2 , \tag{1}$$

where $\mathbb{E}[\cdot]$ represents the unconditional expectation operator. Under the assumption that we can represent the policies $X_{i,t}$ as standardized Ornstein-Uhlenbeck processes, it follows that the dynamics of v_t follow a strictly positive square root process.[12] This policy uncertainty process v_t will affect the real side of the economy through its impact on productivity as well as on output growth.

[11] For the yield effects we refer to, e.g., Kuttner (2001), Piazzesi (2005), Fleming and Piazzesi (2005), Gürkaynak, Sack, and Swanson (2005a), and Wright (2012). For the volatility effects see, e.g., Balduzzi, Elton, and Green (2001), Piazzesi (2005), and de Goeij and Marquering (2006), among others.

[12] For a derivation of this result, see Chapter 6 in Jeanblanc, Yor, and Chesney (2009).

Assumption 1 (Process dynamics)**.** *Productivity A_t, output growth dY_t/Y_t, and government policy uncertainty v_t follow the dynamics*

$$dA_t = (\kappa_A(\theta_A - A_t) + \lambda v_t)\, dt + \sigma_A \sqrt{k_A + v_t} dW_t^A, \tag{2}$$

$$\frac{dY_t}{Y_t} = (\mu_Y + q_A A_t)dt + \sigma_Y \sqrt{k_Y + v_t} dW_t^Y, \tag{3}$$

$$dv_t = \kappa_v \left(\theta_v - v_t\right) dt + \sigma_v \sqrt{v_t} dW_t^v, \tag{4}$$

where $(\theta_A, q_A, \mu_Y) \in \mathbb{R}$ and $(\kappa_A, \sigma_A, \theta_v, \kappa_v, \sigma_v, k_A, k_Y) \in \mathbb{R}^+$. We assume that the correlation between changes in productivity and uncertainty is given by

$$dW_t^A dW_t^v = \rho_{Av} \sqrt{\frac{v_t}{k_A + v_t}} dt, \quad \rho_{Av} \in [-1, 1].$$

All other Brownian motions are mutually independent.

In Equation (2), the parameter $\lambda \in \mathbb{R}$ measures the effect of government policy uncertainty v_t on the growth rate of productivity A_t. Although there is no need to a priori assume a sign restriction on λ, several papers have found it to be negative (see for instance Bloom (2009), Baker, Bloom, and Davis (2016) or Croce, Kung, Nguyen, and Schmid (2012)). The economic rational for this result is fairly intuitive. Higher fiscal uncertainty reduces investment which negatively affects productivity. Running a univariate regression of total factor productivity (TFP) on policy uncertainty, we find that the relationship is indeed negative and highly significant (see Figure 3).

[Figure 3 about here.]

In our setup, government policy uncertainty not only has a first order impact on productivity, it also affects its riskiness. By further inspection of Equation (2), policy uncertainty renders productivity shocks time-varying, i.e., it introduces stochastic volatility to the productivity process. In the extreme case where v_t approaches zero, the diffusive component becomes constant but does not vanish entirely. Along the same line of argumentation, riskiness of output growth is mainly driven by government policy uncertainty shocks.

Having discussed the real side of the economy, we now turn to the specification of the preferences of our representative agent, who is producing a single good at a constant return-to-scale production technology, taking the impact of the government as given. The agent either consumes the good or reinvests it. As in Buraschi and Jiltsov (2005), real monetary holdings M_t^d provide a transaction service by reducing the total amount of gross resources required for a given level of consumption C_t.

Assumption 2 (Preferences of representative agent)**.** *Let $U(X_t)$ denote utility over the real net consumption holdings X and let $\beta > 0$ denote the subjective discount factor. The agent has constant relative risk aversion (CRRA) preferences and maximizes expected utility,*

$$E_t \left[\int_t^{\infty} e^{-\beta s} U(X_s) ds \right], \tag{5}$$

where

$$U(X_t) = \begin{cases} \frac{1}{\gamma}\left(X_t^{\gamma} - 1\right), & \text{if } \gamma < 1, \gamma \neq 0, \\ \log(C_t) + \xi \log\left(M_t^d\right), & \text{if } \gamma = 0, \end{cases} \tag{6}$$

where γ is equal to one minus the coefficient of relative risk aversion. In addition, the real net consumption holdings depend on both consumption C_t and real cash balances M_t^d:[13]

$$X_t = C_t \cdot (M_t^d)^{\xi}, \; 0 \leq \xi \leq 1. \tag{7}$$

In the above specification, utility is additive separable in consumption and real cash-balances whenever $\gamma = 0$.[14] For the general case, however, separability is lost. We now proceed with the formulation of the agent's capital budget constraint.

Assumption 3 (Capital budget constraint)**.** *The real return on capital that can either be allocated to consumption $C_t dt$ or cash balances $M_t^d dt$ or reinvested dK_t is given by*

$$\frac{dK_t}{K_t} + \left(\frac{C_t}{K_t} + \frac{M_t^d}{K_t} \right) dt = \frac{dY_t}{Y_t} - \delta dt, \tag{8}$$

where K_t and Y_t denote the capital accumulation and output process and $\delta \in [0,1]$ is the deprecation rate.

[13] If $\xi = 0$, money does not provide any service and $\xi = 1$ implies that the agent needs to hold exactly one unit of currency for every unit of consumption holdings. Since $0 \leq \xi \leq 1$, a higher level of monetary holdings provides a higher level of transaction services, but at a decreasing return to scale.

[14] Note that we depart from previous literature in that we impose non-separable CRRA utility for our representative agent. The non-separability of preferences helps us to better match simultaneously key statistical moments of yields, their volatilities, and bond excess returns.

Substituting output growth from Equation(3), we obtain the capital accumulation process as

$$\frac{dK_t}{K_t} = -\left(\frac{C_t}{K_t} + \frac{M_t^d}{K_t}\right) dt + \left(\mu_Y + q_A A_t - \delta\right) dt + \sigma_Y \sqrt{k_Y + v_t} dW_t^Y. \tag{9}$$

The capital accumulation process is decreasing in the optimal control variables consumption C_t and money demand M_t^d. This result is intuitive, since higher C_t and M_t^d diminish available resources to be invested in the production technology K_t. Similar to real output, capital is nonstationary whenever it is time-varying in productivity A_t and $\mu_Y \neq \delta$. Furthermore, Equation (9) implies money-neutrality, i.e., monetary shocks do not have an effect on the real side of the economy.[15]

For the monetary sector, we assume that there exists a central bank controlling money supply M_t^S on the basis of a Taylor rule. The monetary authority targets a long term nominal time-varying money growth rate m_t, a capital growth rate $\bar{k}$, and an inflation rate equal to $\bar{\pi}$. We assume that transitory deviations from the optimal long-run money growth exhibit stochastic volatility, which is driven by government policy uncertainty.[16]

Assumption 4 (Monetary sector)**.** *The central bank controls money supply growth according to a Taylor rule as follows:*

$$\frac{dM_t^S}{M_t^S} = m_t dt + \eta_1 \left(\frac{dK_t}{K_t} - \bar{k} dt\right) + \eta_2 \left(\frac{dp_t}{p_t} - \bar{\pi} dt\right) + \sigma_M \sqrt{k_M + v_t} dW_t^M, \tag{10}$$

$$dm_t = \kappa_m(\theta_m - m_t)dt + \sigma_m \sqrt{k_m + v_t} dW_t^m, \tag{11}$$

where p_t and K_t are the price level and capital accumulation processes. We fix the correlation

[15] Whether or not real output and capital are money-shock-neutral is debated in macroeconomics for a long time. The neo-classical Keynesian literature argues that any increase in money supply has to be offset by an equivalent proportional rise in prices and wages. A recent paper using a similar setup as ours is Ulrich (2013), who sticks to this neo-classical view. However, there are a number of reasons why inflation may affect the real economy. See, e.g., Fisher and Modigliani (1978), who argue that inflation has a direct influence on purchasing power, because many private contracts are not indexed. A first quantitative study that allows for dependence of the expected return on capital on inflation is Pennacchi (1991), who uses survey data to identify inflationary expectations. Another channel through which inflation can affect the real economy is through taxation of nominal asset returns. This channel was exploited by Buraschi and Jiltsov (2005) to account for the violation of the expectation hypothesis and the determination of the inflation risk premium. Since policy uncertainty affects the real and nominal side of the economy (through the endogenous equilibrium price level), we assume money-neutrality throughout the paper. However, we acknowledge that a feasible extension of our model is to let the capital accumulation process be a function of the price level. We leave this as an interesting theoretical idea which is worthwhile to be considered.

[16] This modeling assumption is motivated by the observation made in Figure 2. Since uncertainty drives monetary policy decisions (the converse implications do not seem to hold in the data), we let our measure of government policy uncertainty v_t to also affect the nominal side of the economy.

structure as follows:

$$dW_t^M dW_t^v = \rho_{Mv}\sqrt{\frac{v_t}{k_M+v_t}}dt, \quad dW_t^m dW_t^v = \rho_{mv}\sqrt{\frac{v_t}{k_m+v_t}}dt,$$
$$dW_t^A dW_t^m = \rho_{Am}\sqrt{\frac{k_m+v_t}{k_A+v_t}}dt, \quad dW_t^M dW_t^m = \rho_{Mm}\sqrt{\frac{k_m+v_t}{k_M+v_t}}dt,$$

while the shocks W_t^M and W_t^m are independent of all other sources of risk in the economy. Furthermore, to guarantee well-defined correlations, we impose the parameter restrictions $k_A > k_m$, $k_M > k_m$, and $\rho_{Mv}, \rho_{mv}, \rho_{Am}, \rho_{Mm} \in [-1, 1]$.

Similar to the real side of the economy, we assume that the process v_t also drives uncertainty in the Taylor rule. The parameters η_1 and η_2 in Assumption 4 determine the sensitivity of money supply growth with respect to deviations of endogenous capital and inflation from their long-run target levels. For example, if output growth is above its target rate $\bar{k}dt$ and provided that $\eta_1 < 0$, the monetary authority shrinks the money supply, causing interest rates to rise and investment activity to slow down. When inflation is below the target rate $\bar{\pi}dt$ and provided that $\eta_2 < 0$, the central bank's response is to increase money supply. If we have $\eta_1 = \eta_2 = 0$, money supply is exogenous and therefore does not react to deviations from long-run capital nor inflation growth rate. Furthermore, given the money supply rule in Equation (10), policy uncertainty v_t renders money supply time-varying.

Having introduced both the real and monetary side of the economy, we next characterize the representative agent's equilibrium.

Definition 1 (Equilibrium capital stock and money holdings)**.** *Under Assumptions 1 to 4, the representative agent's equilibrium is defined as the vector of optimal consumption, money demand, capital, and price processes $[C_t^*, M_t^{d*}, K_t^*, p_t^*]$ solving the following dynamic Hamilton-Jacobi-Bellmann programming problem*[17]

$$0 = \frac{\partial V(t, K_t, A_t, v_t)}{\partial t} + \max_{\{C_t, M_t^d\}} \left\{ U(C_t, M_t^d) + \mathcal{A}V(t, K_t, A_t, v_t) \right\}, \tag{12}$$

subject to money market-clearing $M_t^S = p_t^* M_t^{d*}$, the bugdet constraint in (8)*, and the transversality condition* $\lim_{t\to\infty} \mathbb{E}\left[e^{-\beta t} V(t, K_t, A_t, v_t)\right] = 0$.

[17] By $\mathcal{A}$ we denote the infinitesimal generator. See, e.g., Øksendal (2003) for technical details.

For the problem in Equation (12), an explicit solution can only be obtained for the log-utility case. The resulting optimal consumption and money demand holdings are proportional to capital K_t. However, up to a first-order approximation in γ, the asymptotic optimal controls C_t and M_t^d remain linear in the state variables K_t, A_t, and v_t. This feature allows us to find an explicit affine representation of the term structure of real and nominal interest rates beyond the log-utility case.[18]

Proposition 2 (Perturbed equilibrium of the representative agent's investment and consumption problem)**.** *In equilibrium, the representative agent's value function is*

$$V(t, K_t, A_t, v_t) = \frac{e^{-\beta t}}{\beta\gamma}\left(\left(e^{\phi(A_t,v_t)}{K_t}^{(1+\xi)}\right)^{\gamma} - 1\right), \tag{13}$$

for some function $\phi(A_t, v_t)$ of the form

$$\phi(A_t, v_t) = \phi_0(A_t, v_t) + \gamma\phi_1(A_t, v_t) + O(\gamma^2). \tag{14}$$

The agent's first order asymptotic optimal controls are

$$C_t^{*,P} = \frac{\beta K_t^*}{1+\xi}\left[1 + \gamma\left(L - \phi_0(A_t, v_t)\right)\right], \qquad M_t^{d*,P} = \xi C_t^*, \tag{15}$$

and the equilibrium first order asymptotic capital accumulation K_t^* and price process p_t^* satisfy

$$\frac{dK_t^*}{K_t^*} = \mu_{K^*}\left(A_t, v_t\right)dt + \sigma_A\sqrt{k_A + v_t}dW_t^Y, \tag{16}$$

$$\begin{aligned}\frac{dp_t^*}{p_t^*} &= \frac{1}{1-\eta_2}\left[\frac{m_t - \eta_1\bar{k} - \eta_2\bar{\pi}}{1-\eta_2} + \frac{\eta_1 - 1}{1-\eta_2}\left(\mu_{K^*}\left(A_t, v_t\right) - \frac{\sigma_Y^2(k_A + v_t)}{(1-\eta_2)}\right)\right]dt \\ &\quad + \frac{\sigma_M\sqrt{k_M + v_t}}{1-\eta_2}dW_t^M + \frac{(\eta_1 - 1)\sigma_Y\sqrt{k_Y + v_t}}{1-\eta_2}dW_t^Y.\end{aligned} \tag{17}$$

where

$$\mu_{K^*}\left(A_t, v_t\right) = \mu_Y + q_A A_t - \beta - \delta + \gamma\beta\left(\phi_0(A_t, v_t) - L\right) \tag{18}$$

denotes the equilibrium drift of the capital accumulation process and $L := \log\left(\frac{\beta^{1+\xi}\xi^{\xi}}{(1+\xi)^{1+\xi}}\right)$ is a constant. Furthermore,

$$\phi_0\left(A_t, v_t\right) = \phi_{00} + \phi_{0A}A_t + \phi_{0v}v_t \tag{19}$$

is an affine function in the state variables (A_t, v_t) with constants ϕ_{00}, ϕ_{0A}, and ϕ_{0v} provided in

[18] Our solution strategy is based on the perturbation method. In particular, we follow the approach of Kogan and Uppal (2001) and approximate our model with respect to the risk aversion parameter around the explicit equilibrium computed under the log-utility assumption. Perturbation methods have been successfully applied in many other studies such as, e.g., Hansen, Heaton, and Li (2008).

Appendix A.2.

Since nominal shocks have no real effects, the equilibrium capital accumulation process is only driven by the real sector of the economy, i.e., by productivity A_t and by the uncertainty process v_t. The weighting factor η_2 on inflation-target deviation enters nonlinearly into the equilibrium price process.[19] We also find that for $\gamma = 0$ the equilibrium capital drift μ_{K^*} becomes independent of v_t. Furthermore, γ has a first-order effect on the representative agent's optimal consumption policy and money demand C_t^* and M_t^*. They both become dependent on policy uncertainty v_t and productivity A_t. Since the constant ϕ_{0A} is positive, whereas ϕ_{0v} is negative (see Equation (A.26)), an increase in productivity increases consumption and money demand, while an increase in uncertainty v_t leads to a decrease. Notably, these effects are absent in the log-utility case.

Proposition 2 also implies that the equilibrium price process is affected by uncertainty shocks through two channels. First, v_t affects the money supply directly through rendering money growth volatility time-varying (see Equation (10)). Secondly, as a direct consequence of the equilibrium market clearing condition $M_t^S = p_t^* M_t^d$ and because the central bank authority is controlling money supply growth based on a Taylor-rule, uncertainty shocks from the real side will propagate to the nominal side whenever $\eta_1 \neq \eta_2 \neq 0$. Hence, by endogenizing money supply growth, we allow for an important link between the real and the nominal sector. This link will prove to be essential in capturing key empirical properties of the yield curve, its corresponding term structure of yield volatility, and bond risk premia.

3 The term structure of nominal interest rates

Having obtained the dynamics of the equilibrium price level, we can now solve for the term structure of nominal and real bond prices. Let $B(t,\tau)$ be the nominal pure discount bond paying one unit of currency at time $T = t+\tau$. The price of the nominal bond must satisfy the following Euler equation

$$B(t,\tau) = e^{-\beta\tau}\mathbb{E}_t\left[\frac{\partial U(C_{t+\tau}^*, M_{t+\tau}^{d*})/\partial C_{t+\tau}^*}{\partial U(C_t^*, M_t^{d*})/\partial C_t^*}\frac{p_t^*}{p_{t+\tau}^*}\right] = e^{-\beta\tau}\mathbb{E}_t\left[\frac{\exp\{-\log\left(K_{t+\tau}^*\right)\}}{\exp\{-\log\left(K_t^*\right)\}}\frac{p_t^*}{p_{t+\tau}^*}\right], \tag{20}$$

[19]Note that for $\eta_2 \approx 1$, small innovations in either A_t or v_t result in dramatic changes in the equilibrium price process. However, from an economic viewpoint, the parameter η_2 should be negative which, as we will see later, is confirmed by the data.

i.e., in equilibrium the investor should be indifferent between consuming one more unit of currency now or investing one unit of currency in the nominal discount bond with time to maturity τ.

Proposition 3 (Equilibrium nominal term structure of interest rates). *In our model economy, the nominal discount bond $B(t,\tau)$ with time to maturity τ is given by*

$$B(t,\tau) = \exp\left[-b_0(\tau) - b_A(\tau)A_t - b_v(\tau)v_t - b_m(\tau)m_t\right], \tag{21}$$

with

$$b_A(\tau) = C_A \frac{1-e^{-\kappa_A \tau}}{\kappa_A}, \tag{22}$$

$$-b_v^{'}(\tau) = Z_{0v}(\tau) + Z_{1v}(\tau)b_v(\tau) + Z_{2v}b_v^2(\tau), \tag{23}$$

$$b_m(\tau) = -\frac{1-e^{-\kappa_m \tau}}{(\eta_2 - 1)\kappa_m}, \tag{24}$$

$$b_0(\tau) = \int_0^\tau C_0(u)du. \tag{25}$$

The constant parameters C_A and Z_{2v} and the deterministic functions $Z_{0v}(\tau)$, $Z_{1v}(\tau)$, and $C_0(\tau)$ are provided in Appendix A.3.

The nominal term structure of interest rates in Proposition 3 belongs to the exponential affine class of term structure models introduced by Duffie and Kan (1996). Consequently, the yield curve $Y(t,\tau)$, defined as

$$Y(t,\tau) = -\frac{1}{\tau}\log\left(B(t,\tau)\right) = \frac{b_0(\tau)}{\tau} + \frac{b_A(\tau)}{\tau}A_t + \frac{b_v(\tau)}{\tau}v_t + \frac{b_m(\tau)}{\tau}m_t, \tag{26}$$

is affine in the state variables A_t, v_t, and m_t. Analyzing the expressions for the constant Z_{2v} and C_A as well as the time-dependent functions Z_{1v}, $Z_{0v}(\tau)$, and $C_0(\tau)$ given in Appendix A.3, we make the following three observations.

First, the target growth rates for output $\bar{k}$ and inflation $\bar{\pi}$, the depreciation rate δ, and output growth rate μ_Y solely affect the intercept of the yield curve, but not its slope. The same holds true for the long-run level of all the three factors, i.e., $\theta_i,\ i = \{A, v, m\}$. In contrast, the central bank's weighting factors η_1 and η_2 affect both the intercept and slope of the yield curve. Moreover, they do so in a nonlinear way.

Our second observation relates to the government policy uncertainty variable v_t. It has a key impact not only on the level of the term structure, but also on its slope. Recall that the trend growth rate of the productivity process A_t not only depends on the long-run level of productivity θ_A, but also on the long-run level of policy uncertainty θ_v. In particular, we have $\mathbb{E}[A_t] = \theta_A + \frac{\lambda\theta_v}{\kappa_A}$. Since $\lambda < 0$, policy uncertainty shocks reduce the long-run level of productivity which has a key impact on the representative agents investment and consumption decision. This effect is amplified if shocks to productivity are persistent, i.e., if κ_A is low. Moreover, this dependence in turn implies that λ also affects the slope of the term structure through $b_v(\tau)$. In summary, fiscal uncertainty affects the term structure of interest rates not only through several channels, but it does so also in a nontrivial way.

As a third observation, we find that the subjective discount factor β and the degree of transaction service money provides ξ also impact the slope of the yield curve through the factor loadings $b_A(\tau)$ and $b_v(\tau)$ whenever $\gamma \neq 0$. Hence, if the representative agent would have log-utility, this channel would be switched off and β and ξ would exhibit only a level effect.

3.1 Equilibrium nominal short rate and bond excess returns

We now discuss how the short end of the term structure of interest rates and the bond risk premium are affected by economic policy uncertainty.

Proposition 4 (Equilibrium nominal short rate and bond risk premium)**.** *In our model economy, we have the following first order asymptotic results:*

1. *The nominal short rate R_t is affine in the state variables:*

$$R_t = K_0 + C_A A_t + C_m m_t + C_v v_t, \tag{27}$$

where the constants K_0, C_A, C_m, C_v are given in Equations (A.48) *to* (A.50).

2. *The bond risk premium $BRP(t,\tau)$ per unit of time is given by*

$$BRP(t,\tau) = \frac{1}{dt}\mathbb{E}_t\left[\frac{dB(t,\tau)}{B(t,\tau)} - R_t dt\right] = b_v(\tau)(k_{0,\Lambda} + k_{1,\Lambda} v_t) =: \Lambda_t^N(\tau), \tag{28}$$

where we denote by $\Lambda_t^N(\tau)$ the nominal risk premium of government policy uncertainty and the constants $k_{0,\Lambda}$ and $k_{1,\Lambda}$ are defined as

$$k_{0,\Lambda} := k_M \frac{\sigma_M \sigma_m \rho_{Mm}}{\eta_2 - 1}, \quad k_{1,\Lambda} := \frac{\sigma_M \sigma_m \rho_{Mm} + \sigma_v \rho_{Mv} \rho_{Mv} \sigma_M}{\eta_2 - 1}. \tag{29}$$

The nominal short rate and the nominal market price of risk are all influenced by both the real and nominal sector of the economy, which is a direct consequence of the Taylor rule in Assumption 4. In the special case when the central bank is inactive, i.e., $\eta_1 = \eta_2 = 0$, it follows from the Taylor equation (11) that money supply is entirely decoupled from the real sector. In this special case, the nominal short rate reduces to

$$R_t = \beta + m_t - (k_M + v_t)\sigma_M^2, \tag{30}$$

which shows that the nominal short rate is only driven by the nominal side of the economy, i.e., output growth and productivity do not affect the nominal short rate. Therefore, introducing an active Taylor rule in (11) opens up a link for real shocks to affect the nominal side of the economy.

In Equation (28) of Proposition 4, bond risk premia are solely driven by policy uncertainty. The loading factor $\Lambda_t^N(\tau)$ allows bond risk premia to be positive or negative depending on the fitted parameter values of the model. As was the case for the nominal short rate, the central bank's reaction parameters η_1 and η_2 affect the nominal risk premium of policy uncertainty through several channels. First, they impact the level of bond risk premia through the term $k_{0,\Lambda}$ in Equation (28). Second, since η_1 and η_2 also enter the factor loading $b_v(\tau)$, they determine the slope and curvature which eventually enables us to match more complex empirical shapes of bond risk premia.

3.2 Data

We obtain monthly Treasury Bill yields with maturity one, two, three, five, seven, and ten years from the Federal Reserve Board ranging from January 1990 until September 2015, from which we bootstrap the zero-coupon yield curve treating the treasury yields as par yields. For our measure of observed volatility, we use realized volatility aggregated on a monthly level from business day data:

$$\mathcal{V}_t(Y(t,\tau)) = \sqrt{100 \times \sum_{d=1}^{D-1} \left(\log\left(\frac{Y(d+1,\tau)}{Y(d,\tau)} \right) \right)^2}, \qquad Y(d,\tau),\ d \in \{1,\ldots,D-1\}, \tag{31}$$

where D denotes the number of daily observations (about 20 business days per month t) and τ is the bond yield maturity. To match $\bar{k}$ we estimate potential output, which we obtain from the database of the Federal Reserve Bank of St. Louis (FRED). The time series is called 'real potential gross domestic product' (GDPPOT).

As a proxy for economic policy uncertainty, our process v_t, we use the economic policy uncertainty (EPU) index constructed by Baker, Bloom, and Davis (2016).[20] The EPU index has three main components, namely a news impact component which is based on news paper discussing economic policy uncertainty, a component that summarizes reports by the Congressional Budget Office (CBO) that compile lists of temporary federal tax code provisions, and a third component called 'economic forecaster disagreement', which draws on the Federal Reserve Bank of Philadelphia's Survey of Professional Forecasters and summarizes data on consumer price forecast dispersion and predictions for purchases of goods and services by state, local and federal government.[21]

To test the robustness of the results from our regression analysis, we introduce different control variates. First, we check whether policy uncertainty has still predictive power after controlling for the state of the economy. The reason why we do so is, arguably, uncertainty about the government's future policy tends to be larger in weaker economic conditions. To proxy for these economic conditions (EC), we use the VIX index, the treasury bond implied volatility (TIV), and the Chicago Fed National Activity Index (CFNAI).[22] Second, we collect two time series, which we refer to as financial variables (FV). They include the monthly log growth rate of the S&P composite dividend yield index (DY), which has been shown to have forecasting power by Fama and French (1989), and, following Campbell and Shiller (1991), the term spread (TS) measured as the ten-year yield less the federal funds rate. Finally, as macroeconomic controls (MC), we collect from Datastream monthly data on industrial

[20]The EPU index can be downloaded from http://www.policyuncertainty.com/. The EPU has been recently used by a number of studies. For instance, Pastor and Veronesi (2013) show that government policy uncertainty carries a risk premium, and that stocks are more volatile and more correlated in times of high uncertainty. Brogaard and Detzel (2015) use the same index and find that economic policy uncertainty forecast future market excess returns. Similarly, Gulen and Ion (2012) show that policy-related uncertainty is negatively correlated with firm and industry level investment. When policy uncertainty increases firm's tend to reduce their investment.

[21]As Kelly, Pástor, and Veronesi (2016) argue, it is difficult isolate exogenous variation in political uncertainty as it likely depends on various factors such as overall macro uncertainty. Therefore, the EPU index may not only capture government related uncertainty, but can be interpreted as a broader measure of uncertainty about economic fundamentals.

[22]The TIV is based on a weighted average of one-month options on treasury bonds with maturity of two, five, ten, and 30 years as a proxy for bond market volatility. It has been widely used in the literature on empirical bond predictability, see for instance Malkhozov, Mueller, Vedolin, and Venter (2016). The CFNAI is obtained from the FRED database.

production (IP) as a measure for real business cycle activity and on the consumer price index (CPI) as a measure for inflation.[23]

3.3 Model Calibration

To calibrate our model to the data, we proceed as follows. In a first step, we perform a maximum likelihood estimation using the EPU index to obtain the parameters of our policy uncertainty process v_t. In a second step, since the Federal Open Market Committee considers an annual inflation of 2% consistent with the Federal Reserve's mandate for price stability and maximum employment, we fix our monthly inflation target $\bar{\pi}$ accordingly. For the central bank's target of long-run output growth, we consider a value for $\bar{k}$ so that it matches the monthly growth rate of potential output (GDPPOT). In a third step, the remaining parameters are then used to calibrate the unconditional model-implied yield and volatility term structure to their empirical unconditional level and volatility term structure. For the yield curve, we average the monthly nominal yields over the whole sample period using

$$\overline{Y}(\tau) = \frac{1}{T}\sum_{t=1}^{T} Y(t,\tau), \quad \tau \in \{1,2,3,5,7,10\}, \tag{32}$$

where T denotes the length of the time series. Our measure for the unconditional realized yield volatility is

$$\overline{\mathcal{V}}(Y(t,\tau)) = \frac{1}{T}\sum_{t=1}^{T-1} \mathcal{V}_t(Y(t,\tau)), \qquad t \in \{1,\dots,T-1\}, \tag{33}$$

with $\mathcal{V}_t(Y(t,\tau))$ given in Equation (31). For calibrating the model, we use as objective function

$$\min_{\Theta} V(\Theta) = \sum_{\tau=1}^{M} \left(\left|\overline{Y}(\tau) - \mathbb{E}[Y(t,\tau)]\right| + \left|\overline{\mathcal{V}}(Y(t,\tau)) - \sqrt{\mathbb{V}[dY(t,\tau)]}\right| \right), \tag{34}$$

where $M = 6$ is the number of maturities and Θ denotes the set of parameters. In total, there are 26 parameters to calibrate. The terms $\mathbb{E}[Y(t,\tau)]$ and $\mathbb{V}[dY(t,\tau)]$ are the model-implied mean and variance of the term structure, which are available in closed form and provided in Appendix A.5.

[Table 1 about here.]

[23] Similar control variables have also been used by Ang and Piazzesi (2003), Evans and Marshall (2007), Ludvigson and Ng (2009), and Joslin, Priebisch, and Singleton (2014) to study the economic determinants of the term structure of nominal interest rates. We also included other commonly used macroeconomic variables, such as employment, personal income, producer price index, the new orders index, and new private housing units started, etc., but our conclusions did not change.

We summarize the calibration results in Table 1. For the policy uncertainty dynamics specified in Assumption 1, we see from Panel A that the speed of mean-reversion κ_v implies a half-life of a shock of $-\log(0.5)/\kappa_v = 4.1$ months. Hence, it takes a little more than four months for a shock to economic policy uncertainty to die out by half. The long-run uncertainty level is characterized by the parameter θ_v, which can also be interpreted (and estimated) as the average economic policy uncertainty level over the estimation time period. Lastly, the parameter σ_v measures the fluctuation of the time series.

From Panel C, it is interesting to note that our estimate of γ implies a relative risk aversion coefficient of 1.47 for the representative agent, which differs substantially from the standard log-utility case with a risk aversion of one.[24] As we have derived in our theoretical analysis, risk-aversion has a first-order effect on the equilibrium capital drift μ_{K^*}, the agent's optimal consumption policy C_t^*, and on money demand M_t^*. Hence, given our estimate of γ, these effects can be substantial and they would be completely absent in a log-utility setting. Furthermore, we see that the calibrated parameter $\xi = 0.463$. Hence, monetary holdings do generate transaction services.

For the monetary sector as specified in Assumption 4, the estimates of the sensitivity parameters η_1 and η_2 turn out to be negative. The central bank is decreasing its money supply whenever capital growth or inflation are above their respective target rates $\bar{k}$ and $\bar{\pi}$. This result confirms our intuition and is in line with the decisions taken by the Federal Reserve. Lastly, the calibrated parameter λ is large and negative. Therefore, policy uncertainty negatively affects both output growth and productivity, corroborating the findings in Rodrik (1991), Bloom (2009), and Gulen and Ion (2012).

[Figure 4 about here.]

[24]Departing from the log utility assumption turns out to have non-trivial effects. After recalibrating the model using the parameters in Table 1 as starting values and restricting $\gamma = 0$ we find the following: While the fit of the average term structure and the volatility curve is comparable, for nearly all maturities, the model fails to produce model-implied beta coefficients, which are statistically indistinguishable from their empirical counterparts. The model-implied betas for the univariate yield regression are now not low enough and do not lie within the 95% confidence band around the empirically estimated coefficients. Furthermore, regarding the model-implied response of volatility to an increase in policy uncertainty, we find that, first, the model-implied coefficients are still positive but too low (not within 95% confidence band) and second, they do not exhibit the hump-shape pattern their empirical counterparts show. Finally, the model-implied beta coefficients underestimate the response of bond excess returns due to an increase in policy uncertainty. This is especially true at the long end.

In Figure 4, we plot the calibrated term structure and its volatility curve for all maturities τ. In Panel A, we observe that the model is able to replicate the overall level, slope, and curvature of the term structure. The model slightly over (under)- estimates the actual level of the yield curve at the short (medium to long) end. However, the overall root mean squared error (RMSE) along the entire term structure is 0.135.[25] As Panel B of Figure 4 shows, our model accurately matches the volatility data for short (one year maturity) to medium term maturities (up to seven years). Importantly, we can replicate the typical volatility hump at two year maturity. The RMSE for the bond volatility calibration is only 0.037.

Many empirical studies find it notoriously difficult to match the term structure of yield volatility, particularly within the framework of affine term structure models. Indeed, even with their more flexible specification of essentially affine price of risk, Buraschi and Jiltsov (2005) acknowledge that their model cannot fully match the second moments of yields. While their model captures the overall declining pattern of volatilities, it fails to generate the volatility hump. Moreover, the model-implied volatilities remain substantially below their empirical counterparts (see their Table 9). Admittedly, they perform quasi-maximum likelihood estimation only on yields. In a more recent endeavor, Malkhozov, Mueller, Vedolin, and Venter (2016) succeed in generating a hump, but they fail to match the level accurately (see their Figure 6). These results exemplify the intrinsic challenge of matching both the level of volatilities and the volatility hump.[26] In contrast, we find that our model is rich enough to match the unconditional term structure and the high level of volatility simultaneously, while replicating the typical hump-shape in the bond volatility curve.

[25] The error is calculated as RMSE $= \sqrt{\frac{1}{T}\sum_{\tau\in M}\left(\mathbb{E}[Y(t,\tau) - \overline{Y}(t,\tau)]\right)^2}$ where $M = 6$ is the number of maturities, $\mathbb{E}[Y(t,\tau)]$ is the model-implied unconditional yield curve and $\overline{Y}(t,\tau_i)$ defines the unconditional sample average of yields. The same formula is applied compute the error along the volatility and correlation curve.

[26] Already Shiller (1979) shows that long term bond yields exhibit excess volatility relative to their model-implied values. From Piazzesi and Schneider (2006) we know that their representative agent-based model explains a smaller fraction of observed volatility of the long-end yields than of the short-end yields. Xiong and Yan (2010) argue that excess bond volatility might be due to differences in beliefs about the long-run level of inflation. They show that a higher belief dispersion leads to volatility amplification which allows them to account not only for the empirically observed high bond yield volatility, but also for the hump-shape of the term structure of bond volatility. Closely related to the 'excess volatility puzzle' phenomenon are also the findings of Gürkaynak, Sack, and Swanson (2005b). They document that bond yields exhibit excess sensitivity to macroeconomic announcements.

3.4 Yield curve and policy uncertainty

Based on our fitted model parameters, we can now explore a series of implications regarding the effect of policy uncertainty on the term structure of interest rates, bond yield curve, and bond risk premia. We start our analysis by investigating the relationship between nominal bond yields and policy uncertainty as measured by the EPU index. The preliminary empirical analysis between nominal bond yields and policy uncertainty shows that there is significant negative correlation between economic policy uncertainty and the level of the yield curve, which we observed in Figure 1, Panel A. Indeed, given the yields in Equation (26), we have that

$$\frac{\partial Y(t,\tau)}{\partial v_t} = \frac{b_v(\tau)}{\tau} < 0, \qquad \forall \tau \geq 0. \tag{35}$$

The inequality in Equation (35) follows from the calibrated factor loadings $b_v(\tau)$, which we report in Table 2. Clearly, the loadings $b_v(\tau)$ are negative for all $\tau \geq 0$. Hence, nominal yields decline when policy uncertainty increases.

[Table 2 about here.]

A second implication of our term structure model is that nominal yields are negatively correlated with economic policy uncertainty, i.e., $\mathrm{Corr}\left[Y(t,\tau), v_t\right] \leq 0$, given the fitted parameters in Table 1. Evaluating the model-implied correlations across different maturities gives an average correlation between nominal yields and the EPU of -0.667, which is comparable with the average sample correlation of -0.4835.

For the impact of a change in policy uncertainty on the term structure of bond yield volatility, we find that the conditional model-implied volatility curve $\sqrt{\mathbb{V}_t[dY(t,\tau)]}$ given in (A.57) is increasing in policy uncertainty v_t:

$$\frac{\partial \sqrt{\mathbb{V}_t[dY(t,\tau)]}}{\partial v_t} = \frac{\Gamma_1(\tau)}{2\sqrt{\Gamma_0(\tau) + \Gamma_1(\tau) v_t}} > 0, \ \forall \tau \geq 0 \tag{36}$$

where the functions $\Gamma_0(\tau),\ \Gamma_1(\tau) : \tau \to \mathbb{R}_+$ are given in Equation (A.59). Furthermore, given the calibrated parameters in Table 1, the expression in Equation (36) is hump-shaped across maturity τ and peaks at two year maturity (see Figure 4).

4 Empirical analysis

Given the calibrated values in Table 1, our term structure model gives rise to several model-implied predictions. From the discussion above, we can formulate four testable hypotheses.

Hypothesis 1 (H1): Nominal yields fall when policy uncertainty increases. This hypothesis follows from Equation (35).

Hypothesis 2 (H2): Higher (lower) economic policy uncertainty increases (decreases) nominal yield volatility. This hypothesis follows from the positive loading function $\Gamma_1(\tau)$ in Equation (36).

Hypothesis 3 (H3): The hump shape of the term structure of bond yield volatility is driven by policy uncertainty and peaks around the two-year maturity. This hypothesis follows from the expression in Equation (36) and its shape, given the calibrated parameters in Table 1.

Hypothesis 4 (H4): Bond risk premia are increasing in policy uncertainty. This hypothesis follows from Equation (28) and the calibrated parameters in Table 1, which imply $\partial \Lambda_t^N(\tau)/\partial v_t > 0$.

To empirically test the above hypotheses, we regress nominal yields (Y), yield volatility (V), and bond excess returns (BRP) on the EPU index. We then compare these estimated regression coefficients to their model-implied counterparts, which we derive in Proposition 5 below.

Proposition 5 (Model-implied regression coefficients)**.** *In our model economy, we have the following results for the model-implied regression coefficients β^{MI}:*

1. *The unconditional model-implied regression coefficient β_Y^{MI} of a univariate regression of nominal yields $Y(t,\tau)$ onto policy uncertainty v_t is given by*

$$\beta_Y^{MI} = \frac{\mathbb{C}[Y(t,\tau), v_t]}{\mathbb{V}(v_t)} = \frac{1}{\tau}\left(\frac{b_A(\tau)\mathbb{C}[A_t, v_t]}{\mathbb{V}(v_t)} + b_v(\tau) + \frac{b_m(\tau)\mathbb{C}[m_t, v_t]}{\mathbb{V}[v_t]}\right). \tag{37}$$

2. *The first-order approximate unconditional model-implied regression coefficient* β_Y^{MI} *of a univariate regression of bond volatility* $\sqrt{\mathbb{V}_t[dY(t,\tau)]}$ *onto policy uncertainty* v_t *is*

$$\beta_V^{MI} = \frac{\mathbb{C}\left(\sqrt{\mathbb{V}_t[dY(t,\tau)]}, v_t\right)}{\mathbb{V}(v_t)} \approx \frac{\Gamma_1(\tau)}{2\sqrt{\Gamma_0(\tau)+\Gamma_1(\tau)\theta_v}}, \tag{38}$$

where the Taylor expansion of $\sqrt{v_t}$ *is evaluated at the unconditional first moment of* v_t*, i.e.,* $\mathbb{E}(v_t) = \theta_v$*.*

3. *The theoretical slope coefficient* β_{BRP}^{MI} *of a univariate regression of excess returns* $r_{t,t+h}^{E,\tau} = \log\left(B(t+h,\tau-h)\right) - \log\left(B(t,\tau)\right) - Y(t,h)$ *over a time period* t *to* $t+h$ *onto policy uncertainty* v_t *is given by*

$$\beta_{BRP}^{MI} = \frac{\mathbb{C}\left(r_{t,t+h}^{E,\tau}, v_t\right)}{\mathbb{V}(v_t)} = \frac{\Omega(t,\tau,h)}{\mathbb{V}(v_t)}, \tag{39}$$

where the covariance term $\Omega(t,\tau,h)$ *is given in Equation* (A.66) *and the denominator is given in Equation* (A.7)*.*

In what follows, we compare the estimated regression coefficients $\widehat{\beta}_i$ with their theoretical counterparts β_i^{MI}, $i \in \{Y, V, BRP\}$, for each univariate regression of yields, bond volatility, and bond excess return onto policy uncertainty. To test for robustness of our results, we subsequently add several control variables.

4.1 Policy uncertainty and the yield curve

To investigate the relationship between the yield curve and the EPU index, we start with the following univariate regression,

$$Y(t,\tau) = \beta_0 + \beta_Y EPU_t + \epsilon_t, \quad \tau = 1, 2, 3, 5, 7, 10, \tag{40}$$

where EPU_t denotes the economic policy uncertainty (EPU) index at time t and ϵ_t is the regression error term.[27] According to hypothesis **H1**, we expect the coefficient in Equation (40) to be negative for every maturity τ. Furthermore, if the model is able to replicate the data, the theoretical beta β_Y^{MI} should lie within the confidence interval around its empirical counterpart $\widehat{\beta}_Y$.

[27] To address potential concerns about robustness of our results, following Newey and West (1994), we compute standard errors with five lags to account for heteroskedasticity and autocorrelation (HAC) in residuals.

[Figure 5 about here.]

Figure 5 presents the results from the comparison of β_Y^{MI} and $\widehat{\beta}_Y$. We can draw two conclusions. First, from Panel A we observe that the impact of an increase in economic policy uncertainty on the level of yields is negative and statistically significant at the 5% level across all maturities, confirming our hypothesis **H1**. Moreover, not only is the EPU index a statistically significant predictor, its impact on the level of the term structure is also economically substantial. For instance, the estimated coefficients imply that a one standard deviation change in the EPU index will lead to a decline in the one year yield of 1.07%.[28] Second, the model-implied regression coefficients β_Y^{MI} lie within the 95% confidence interval. Hence, the model-implied betas are not statistically different from their empirical regression betas $\widehat{\beta}_Y$.

To check for robustness, we add different controls to the regression equation, controlling for economic, financial, and macroeconomic conditions, as discussed in Section 3.2. Panel B of Figure 5 plots the resulting regression coefficient $\widehat{\beta}_Y$ together with the confidence bounds. We find that the negative relationship between nominal yields and policy uncertainty is robust using different controls and remains statistically significant. Table 3 provides an overview of the different regression results. Most strikingly, the EPU is significant for any maturity and across all regressions. The predictive power of the EPU index is reflected also in the R^2_{adj} values. Using the EPU as single predictor, it explains 15% to 28% of the variation in the term structure, depending on the time to maturity. Moreover, whereas the R^2_{adj} values essentially stay the same after adding different controls for economic and macroeconomic conditions, they increase considerably when adding the financial variables. Both the S&P dividend yield and the term spread are statistically significant. Adding macroeconomic controls to the regression equation has only a marginal impact on the statistical significance and the magnitude of the EPU index, and leads only to a small increase in the R^2_{adj} values.[29]

[Table 3 about here.]

[28] This estimate is computed as the estimated slope coefficient $\widehat{\beta}_{Y,1} = -3.715$ for the one year yield times the standard deviation of the EPU (0.34).

[29] The increase in R^2_{adj} is entirely driven by inflation as it has a statistically significant and positive impact on the term structure of nominal interest rates.

4.2 The term structure of bond yield volatility and policy uncertainty

Hypotheses **H2** and **H3** state that the inclusion of a time-varying policy risk factor not only raises the level of the bond volatility curve, but is also a key driver in generating the empirically observed hump shape of the bond volatility term structure. Thus, we should observe a positive regression coefficient peaking around the two year maturity bucket, similar to the realized bond volatility curve in Figure 4. To test these predictions, we regress the conditional volatility $\mathcal{V}_t(Y(t,\tau))$ on the EPU index:

$$\mathcal{V}_t(Y(t,\tau)) = \beta_0 + \beta_V EPU_t + \epsilon_t, \quad \tau = 1, 2, 3, 5, 7, 10. \tag{41}$$

To check for robustness, we include the same control variables as in the previous section.

[Figure 6 about here.]

Figure 6, Panel A, corroborates our hypotheses. The regression coefficient is positive and statistically significant at the 5% level across all maturities. Hence, in line with hypothesis **H2**, an increase in economic policy uncertainty raises the level of bond yield volatility. Moreover, all model-implied regression coefficients $\beta^{MI}_{V,\tau}$ lie within the 95% confidence interval. Hence, our model-implied betas are statistically not different from their empirical counterparts $\widehat{\beta}_{V,\tau}$. Remarkably, both the model-implied and the empirical estimate peak at two year maturity, thereby confirming our hypothesis **H3** that the impact of policy uncertainty is hump-shaped in time to maturity. The impact of the EPU on bond yield volatilities is not only statistically but also economically significant. At the two year maturity, a one standard deviation increase in the EPU increases the monthly realized bond volatility by 0.425%.[30] Comparing this number to the average level of the bond volatility curve as shown in Panel B of Figure 4, where the unconditional level of the two year yield volatility is 1.24%, the estimated impact of a one standard deviation increase in the EPU corresponds to a 34% increase in the two year bond yield volatility.

In Panel B of Figure 6, we plot the regression coefficient when we include all the variables controlling for economic, financial, and macroeconomic conditions, as we did in Figure 5. Also in this case, our results are robust. The impact of economic policy uncertainty on bond yield volatilities

[30] This number is computed as the estimated slope coefficient times the standard deviation of EPU index (0.21).

remains positive and hump-shaped. Table 4 gives an overview of the different regression results. As for the yield regressions, the EPU remains a highly significant factor in determining bond yield volatilities. Our hypotheses **H2** and **H3** are supported by the data even after controlling for a number of variables. Not only does realized volatility rise when policy uncertainty increases, its effect is also hump-shaped with a peak around a maturity of two years.[31] Moreover, the EPU index exhibits high explanatory power. Across different maturities, the R^2_{adj} values range between 34% and 45%.

[Table 4 about here.]

4.3 Bond risk premia

To explore the predictive power of the EPU index for future bond excess returns, we run the univariate regression

$$r^{E,\tau}_{t,t+1} = \beta_0 + \beta_{BRP} EPU_t + \epsilon_t, \quad \tau = 1, 2, 3, 5, 7, 10, \tag{42}$$

where the bond excess return $r^{E,\tau}_{t,t+1}$ is defined as the log-excess return from buying at time t a bond with time-to-maturity τ and selling it after one month:

$$r^{E,\tau}_{t,t+1} = \log\left(B(t+1, \tau-1)\right) - \log\left(B(t,\tau)\right) - Y(t,1). \tag{43}$$

Our Hypothesis **H4** states that government policy uncertainty should explain bond risk premia and, given the calibrated parameters in Table 1, that bond excess returns are increasing in policy uncertainty. Therefore, we expect the estimated coefficient $\widehat{\beta}_{BRP}$ in the regression (42) to be positive and increasing in maturity.

[Figure 7 about here.]

Figure 7, Panel A, plots the theoretical betas β^{MI}_{BRP} together with their empirical counterparts $\widehat{\beta}_{BRP}$ resulting from regression (42). In general, our model predictions **H4** are confirmed. Not only is

[31] An exception is when we add only the controls for the economic conditions. In this case, the hump-shaped impact of the EPU index on the term structure disappears. However, the impact of the economic condition controls is insignificant.

the EPU index a statistically significant predictor, its impact on bond risk premia is also economically substantial. For instance, the estimated coefficients imply that a one standard deviation change in the EPU index will lead to an increase of 2.39% in the expected 10-year bond excess return, which is again substantial given the average 10-year bond excess return of 5.391%.[32] Furthermore, except for the 2 and 3 years excess return, the null hypothesis that the model-implied beta coefficients are equal to their empirical counterparts is not rejected at the 5% confidence level.

To check the robustness of our results, we enrich our previous set of controls for economic, financial, and macroeconomic conditions with an additional set of control variables extracted from the yield curve. The literature on bond risk premia usually tests the predictive power of a new predictor variable against the routinely used Cochrane and Piazzesi (2005) factor (CP), which is constructed based on a tent-shaped linear combination of forward rates.[33] To construct the CP factor, we use the log forward rates, defined as $F(t,\tau) = \log(B(t,\tau-1)/B(t,\tau))$. Furthermore, from the covariance matrix of yields, we extract the first three principal components which are commonly referred to as 'level', 'slope', and 'curvature' (see Litterman and Scheinkman (1991)).

[Table 5 about here.]

From Figure 7, Panel B, the impact of the EPU on bond risk premia changes only slightly. While we still observe a substantial positive relationship between the EPU and bond risk premia that increases with maturity, some of the statistical significance is lost at the short end of the curve. Table 5 gives a detailed overview on our bond risk premium regression. We observe that, when adding the principal components and the CP factor, the impact of the EPU index is significantly reduced. The CP factor is highly significant for most maturities. Moreover, the explanatory power increases substantially. While the R^2_{adj} values for the univariate regression on the EPU range between 1% for the two-year maturity to 32% for the 10 year maturity, the R^2_{adj} ranges from 31% (two years) to 72% (five years) when including the term structure controls. Most of this increase is driven by the CP

[32] This estimate is computed as the estimated slope coefficient $\widehat{\beta}_{BRP,10} = 6.746$ for the 10 year excess return times the standard deviation of the EPU (0.80).

[33] A detailed description of the construction of this factor is given in Cochrane and Piazzesi (2005). To avoid collinearity problems, we only include the current one year yield $Y(t,1)$ and the five and ten year forward rates and we do not restrict the regression coefficients to sum up to one.

factor. Adding the controls for economic, financial, and macroeconomic condition, we observe that the S&P dividend yield and the current slope of the term structure are highly significant, leading to a sharp increase in R^2_{adj} values at shorter maturities. Lastly, adding the macro control variables leaves the explanatory power of the EPU index essentially unchanged. Hence, overall our hypothesis **H4** is confirmed in that economic policy uncertainty has a positive and, except at the short end, significant impact on bond risk premia.

5 Conclusion

In this paper, we present a tractable dynamic general equilibrium model that allows us to study the impact of policy uncertainty shocks on the term structure of interest rates, the yield volatility curve, and bond risk premia. Unlike previous literature, we equip the representative agent with non-separable CRRA preferences and derive nominal bond prices in closed form up to a first-order perturbation in the agent's risk aversion coefficient.

Even though our model belongs to the class of affine term structure models, it is capable of reproducing simultaneously the shape of the term structure of yields and volatilities, including the hump-shape of volatilities which several affine term structure models fail to achieve. Moreover, our model leads to a set of predictions for policy uncertainty and its impact on the interest rate term structure. Our empirical tests provide strong support for these predictions. First, our calibrated model implies a negative relationship between policy uncertainty and the level of yields, which is in line with the well known empirical flight-to-quality behavior, i.e., investors seek safe assets in times of high uncertainty. Second, our model is able to match the empirically observed (hump-shaped) increase in yield volatility in response to uncertainty shocks. Lastly, our model predicts that the risk premia on policy uncertainty is positive and increasing, which is also largely supported by the data.

A Proofs

A.1 Moment Formulas

Let $0 \leq t \leq s$ and define $v_s = e^{-\kappa_v(s-t)}V_s,\ V_t = v_t > 0$. Then an application of Itô's lemma shows that

$$dv_s = -\kappa_v e^{-\kappa_v(s-t)}V_s ds + e^{-\kappa_v(s-t)}dV_s = -\kappa_v v_s ds + e^{-\kappa_v(s-t)}dV_s \tag{A.1}$$

Rearranging this expression, we find

$$e^{-\kappa_v(s-t)}dV_s = \kappa_v\theta_v v_s ds + \sigma_v\sqrt{v_s}dW_s^v, \tag{A.2}$$

which after using $v_s = e^{-\kappa_v(s-t)}V_s$ gives us

$$v_s = v_t e^{-\kappa_v(s-t)} + \theta_v\left(1 - e^{-\kappa_v(s-t)}\right) + \sigma_v e^{-\kappa_v(s-t)}\int_t^s e^{\kappa_v(u-t)}\sqrt{v_u}dW_u^v. \tag{A.3}$$

Using the same substitution, we get for the productivity A_s and money supply factor m_s the following expressions

$$\begin{aligned}
A_s &= A_t e^{-\kappa_A(s-t)} + \theta_A\left(1 - e^{-\kappa_A(s-t)}\right) + \lambda e^{-\kappa_A(s-t)}\int_t^s e^{\kappa_A(u-t)}v_u du \\
&\quad + \sigma_A e^{-\kappa_A(s-t)}\int_t^s e^{\kappa_A(u-t)}\sqrt{k_A + v_u}dW_u^A,\ A_t \in \mathbb{R} \qquad \text{(A.4)} \\
m_s &= m_t e^{-\kappa_m(s-t)} + \theta_m\left(1 - e^{-\kappa_m(s-t)}\right) + \sigma_m e^{-\kappa_m(s-t)}\int_t^s e^{\kappa_m(u-t)}\sqrt{k_m + v_u}dW_u^m,\ m_t \in \mathbb{R}_+
\end{aligned}$$

The unconditional covariance of v_t and v_s,

$$\mathbb{C}(v_s, v_t) = \mathbb{E}[v_s v_t] - \mathbb{E}[v_s]\mathbb{E}[v_t] = \mathbb{E}[\mathbb{E}_t[v_s]v_t] - \mathbb{E}[v_s]\mathbb{E}[v_t], \tag{A.5}$$

is then obtained as

$$\mathbb{C}(v_s, v_t) = \mathbb{E}\left[\left(\theta_v + (v_t - \theta_v)e^{-\kappa_v(s-t)}\right)v_t\right] - \theta_v^2 = \left(\mathbb{E}\left[v_t^2\right] - \theta_v^2\right)e^{-\kappa_v(s-t)} = \frac{\theta_v\sigma_v^2}{2\kappa_v}e^{-\kappa_v(s-t)}, \tag{A.6}$$

with $0 \leq t \leq s$. The unconditional variance follows directly as

$$\mathbb{V}[v_t] = \frac{\theta_v\sigma_v^2}{2\kappa_v}. \tag{A.7}$$

Similarly, for the unconditional covariance between the productivity process A_s, monetary factor m_s, and the policy uncertainty process v_t at time t and s, $0 \leq t \leq s$, we get

$$\mathbb{C}(A_s, v_t) = \frac{e^{-(s-t)\kappa_A}\theta_v\rho_{Av}\sigma_v\sigma_A}{\kappa_A+\kappa_v} + \frac{e^{-(s-t)\kappa_A}\theta_v\lambda\sigma_v^2}{2\kappa_v(\kappa_A+\kappa_v)} + \frac{\left(e^{-s\kappa_v} - e^{-(s-t)\kappa_A - t\kappa_v}\right)\theta_v\lambda\sigma_v^2}{2\kappa_v(\kappa_A-\kappa_v)}, \tag{A.8}$$

$$\mathbb{C}(m_s, v_t) = \frac{\rho_{mv}\sigma_m\sigma_v}{\kappa_m+\kappa_v}e^{-\kappa_m(s-t)}. \tag{A.9}$$

The expressions for $\mathbb{C}[A_t, m_s]$ and $\mathbb{V}[A_t]$ can be derived along the same lines.

A.2 Proof of Proposition 2

The optimal consumption and investment problem is

$$\max_{C_s, M_s^d} \mathbb{E}_t\left[\int_t^\infty e^{-\beta s}U\left(C_s, M_s^d\right)ds\right], \tag{A.10}$$

where $U(C_t, M_t^d) = \frac{1}{\gamma}\left(\left(C_t(M_t^d)^\xi\right)^\gamma - 1\right)$ subject to the capital constraint in Equation (8). To simplify our notation, let $X_t = (A_t, v_t)$ such that the value function is given by

$$V = V\left(t, K_t, X_t\right) = \max_{\{C_s, M_s^d\}_{t\leq s<\infty}} \mathbb{E}_t\left[\int_t^\infty e^{-\beta s}U\left(C_s, M_s^d\right)ds\right]. \tag{A.11}$$

In equilibrium, there exists a value function $V\left(t, K_t, X_t\right)$ and control variables $\{C_t, M_t^d\}$ satisfying the HJB equation

$$-\frac{\partial V\left(t, K_t, X_t\right)}{\partial t} = \max_{\{C_t, M_t^d\}}\left\{a(t)U(C_t, M_t^d) + \mathcal{A}V\left(t, K_t, X_t\right)\right\}, \tag{A.12}$$

where $a(t) = e^{-\beta t}$. By standard time-homogeneity arguments for infinite horizon problems we have that

$$\begin{aligned}
e^{\beta t}V\left(t, K_t, X_t\right) &= \max_{\{C_s, M_s^d\}_{t\leq s<\infty}} \mathbb{E}_t\left[\int_t^\infty e^{-\beta(s-t)}U\left(C_s, M_s^d\right)ds\right] \\
&= \max_{\{C_{t+u}, M_{t+u}^d\}_{t\leq u<\infty}} \mathbb{E}_t\left[\int_t^\infty e^{-\beta u}U(C_{t+u}, M_{t+u}^d)du\right] \\
&= \max_{\{C_u, M_u^d\}_{0\leq u<\infty}} \mathbb{E}_0\left[\int_0^\infty e^{-\beta u}U(C_u, M_u^d)du\right] \\
&\equiv H\left(K_t, X_t\right),
\end{aligned} \tag{A.13}$$

where the third equality follows because the optimal robust control is Markov and $H(K_t, X_t)$ is independent of time. Therefore, we conjecture that the value function has the following form

$$V(t, K_t, X_t) = a(t)H(K_t, X_t) = e^{-\beta t} H(K_t, X_t) = \frac{e^{-\beta t}}{\gamma}\left(\left(e^{\phi(X_t)} {K_t}^{\xi}\right)^{\gamma} - 1\right), \tag{A.14}$$

where $\phi : \mathbb{R}^2 \to \mathbb{R}$ is $\mathcal{C}^N$-differentiable function of the state vector X_t that needs to be determined in equilibrium. Inserting (A.14) into the HJB equation in (A.12), the first order conditions are

$$C_t^* = \frac{K_t\left({K_t}^Q e^{\phi(X_t)}\right)^{-\gamma}}{\beta Q}\left(\frac{\beta Q\left({K_t}^Q e^{\phi(X_t)}\right)^{\gamma}\left(\left(-\frac{(\gamma-1)\beta^{\frac{1}{1-\gamma}} Q^{\frac{1}{1-\gamma}} {K_t}^{\frac{1-\gamma Q}{\gamma-1}} e^{\frac{\gamma\phi(X_t)}{1-\gamma}}}{\xi}\right)^{\frac{1-\gamma}{\gamma\xi+\gamma-1}}\right)^{-\xi}}{K_t}\right)^{\frac{\gamma}{\gamma-1}}, \tag{A.15}$$

$$M_t^{d*} = \left(\frac{(1-\gamma)\beta^{\frac{1}{1-\gamma}} Q^{\frac{1}{1-\gamma}} {K_t}^{\frac{1-\gamma Q}{\gamma-1}} e^{\frac{\gamma\phi(X_t)}{1-\gamma}}}{\xi}\right)^{\frac{1-\gamma}{\gamma\xi+\gamma-1}}, \tag{A.16}$$

with $Q = 1 + \xi$. In general, the function $\phi(X_t)$ cannot be obtained in closed form. However, we can obtain an asymptotic expansion to $\phi(X_t)$ with respect to the risk aversion parameter γ. Assuming a power series expression for $\phi(X_t)$ in γ as follows

$$\phi(X_t) = \phi_0(X_t) + \gamma\phi_1(X_t) + O(\gamma^2),$$

where $\phi_0(X, t)$ is obtained from the logarithmic utility case, we can then solve the HJB problem in Equation (A.12) for $\gamma \neq 0$ in closed form. To do so, we solve first the HJB problem in the case where utility of the representative agent is logarithmic.

A.2.1 Log-utility case

Note that for $\gamma \to 0$ the utility reduces to

$$\lim_{\gamma\to 0} U(C_t, M_t^d) = \log(C_t) + \xi\log\left(M_t^d\right). \tag{A.17}$$

The optimal consumption and investment problem is[34]

$$\max_{C_s, M_s^d} \mathbb{E}_t \left[\int_t^{\infty} e^{-\beta s} \left[\log(C_s) + \xi \log \left(M_s^d \right) \right] ds \right], \tag{A.18}$$

subject to the capital constraint in Equation (8). In equilibrium, there exists a value function $V^{log}(t, K_t, X_t) = a(t)H^{log}(K_t, X_t)$ and control variables $\{C_t, M_t^d\}$ satisfying the HJB equation

$$-\frac{\partial a(t)H^{log}(K_t, X_t)}{\partial t} = a(t) \left(\max_{\{C_t, M_t^d\}} \left\{ U^{log}(C_t, M_t^d) + \mathcal{A}H^{log}(K_t, X_t) \right\} \right). \tag{A.19}$$

We consider the following linear conjecture for the value function $V^{log}(\cdot)$:

$$a(t)H^{log}(K_t, X_t) = e^{-\beta t} \left[Q \log(K_t) + \phi_0(X_t) \right], \tag{A.20}$$

where the function $\phi_0(X_t)$ is affine in the state variables, i.e.,[35]

$$\phi(X_t) = \phi_{00} + \phi_{0A} A_t + \phi_{0v} v_t. \tag{A.21}$$

Applying the generator to Equation (A.20), using the productivity, the policy uncertainty, and the capital accumulation dynamics in Equation (2), (4) and (9), we obtain

$$\begin{aligned} \mathcal{A}V^{log} &= \frac{\partial V^{log}}{\partial K} \mu_K(A_t) + \frac{\partial V^{log}}{\partial A} \mu_A(A_t, v_t) + \frac{\partial V^{log}}{\partial v} \mu_v(v_t) + \frac{1}{2} \frac{\partial^2 V^{log}}{\partial K^2} \sigma_K^2(v_t) \\ &= Q \left[-\left(\frac{C_t}{K_t} + \frac{M_t^d}{K_t} \right) + (\mu_y + q_A A_t - \delta) - \frac{1}{2} \sigma_Y^2 (k_Y + v_t) \right] \qquad \text{(A.22)} \\ &+ (\kappa_A (\theta_A - A_t) + \lambda v_t) \phi_{0A} + \kappa_v (\theta_v - v_t) \phi_{0v}. \qquad \text{(A.23)} \end{aligned}$$

The first order optimality conditions for consumption and money holdings are

$$\frac{e^{-\beta t}}{C_t} - \frac{Q}{\beta} \frac{e^{-\beta t}}{K_t} = 0 \Longleftrightarrow C_t^* = \frac{\beta K_t}{Q}, \tag{A.24}$$

$$\frac{e^{-\beta t} \xi}{M_t^d} - \frac{Q}{\beta} \frac{e^{-\beta t}}{K_t} = 0 \Longleftrightarrow M_t^{d*} = \frac{\beta \xi K_t}{Q}. \tag{A.25}$$

Substituting the optimal controls C_t^* and M_t^{d*} into Equation (A.19) and matching coefficients of $\log(K_t)$, A_t, v_t, and the constant terms, we obtain

$$Q = 1 + \xi, \quad \phi_{0A} = \frac{q_A Q}{(\kappa_A + \beta)}, \quad \phi_{0v} = \frac{2\lambda \phi_{0A} - Q \sigma_Y^2}{2(\kappa_v + \beta)}, \tag{A.26}$$

[34] The proof of this proposition is similar to the one presented in Buraschi and Jiltsov (2005).

[35] Due to money neutrality, the nominal factor m_t does not affect the real side of the economy. Because of this property, the affine conjecture in (A.21) does not include m_t since for optimality it has to be the case that $\phi_{0m} = 0$.

and ϕ_{00} is a lengthy expression depending only on the model parameters. The coefficients are all uniquely determined, state-independent and also independent of K_t. Substituting the expressions back into the HJB equations verifies our guess in (A.20). Furthermore, to prove the transversality condition, we need to verify that $\lim_{t\to\infty} \mathbb{E}\left[e^{-\beta t} H(K_t, A_t, v_t)\right] = 0$. Since the value function guess is additive in $\log(K_t)$ as well as A_t and v_t, we can treat each term individually. First, since both A_t and v_t are mean reverting processes, multiplying the guess by $e^{-\beta t}$ will always converge to zero exponentially fast as long as $\beta > 0$. Next, applying Itô's formula to $Z_t := \log(K_t)$ gives

$$dZ_t = \left(\mu_{K^*}(A_t, v_t) - \frac{1}{2}\sigma_Y^2(k_Y + v_t)\right) dt + \sigma_Y \sqrt{k_Y + v_t} dW_t^Y, \tag{A.27}$$

which after integrating from 0 to t and taking unconditional expectation gives

$$\mathbb{E}[Z_t] = Z_0 + \left(\mu_{K^*}\left(\theta_A + \frac{\lambda\theta_v}{\kappa_A}, \theta_v\right) - \frac{1}{2}\sigma_Y^2(k_Y + \theta_v)\right) t, \quad Z_0 \in \mathbb{R}_+. \tag{A.28}$$

where $\mu_{K^*}\left(\theta_A + \frac{\lambda\theta_v}{\kappa_A}, \theta_v\right)$ refers to the drift of capital in (18) evaluated at the unconditional first moments of productivity A_t and policy uncertainty v_t. Finally, pre-multiplying (A.28) by $e^{-\beta t}$ and passing to the limit using l'Hôpital's rule shows that

$$\lim_{t\to\infty} = e^{-\beta t}\mathbb{E}[Z_t] = 0,$$

which verifies that the transversality condition is satisfied.

A.2.2 Perturbed solution

Let $V = V(t, K_t, X_t)$ denote the value function as given in Equation (A.11). Utility is now given by the following non-separable preferences

$$U(C_t, M_t^d) = \frac{1}{\gamma}\left(\left(C_t(M_t^d)^\xi\right)^\gamma - 1\right). \tag{A.29}$$

From the HJB equation in (A.12), the optimal consumption C_t^* and money demand M_t^{d*} policy holdings are given by[36]

$$C_t^* = \left(M_t^d\right)^{-\xi} \left(\frac{Q\left(M_t^d\right)^{-\xi}\left({K_t}^Q e^{\phi(X_t)}\right)^{\gamma}}{\beta K_t}\right)^{\frac{1}{\gamma-1}}, \tag{A.30}$$

$$M_t^{d*} = \left(\frac{Q C_t^{-\gamma} {K_t}^{\gamma Q-1} e^{\gamma\phi(X_t)}}{\beta\xi}\right)^{\frac{1}{\gamma\xi-1}}, \tag{A.31}$$

where $Q = 1+\xi$. Next, inserting optimal money demand (A.31) into the first order condition of consumption (A.30), using the power series representation of $\phi(X_t)$ as given in Equation (14) and perturbing the resulting expression around the log-utility case (and analogously for optimal money demand), the perturbed optimal consumption and money holdings are given by

$$C_t^{*,P} = \frac{\beta K_t}{(1+\xi)}\left[1+\gamma\left(\log\left(\frac{\beta^{1+\xi}\xi^{\xi}}{(1+\xi)^{1+\xi}}\right)-\phi_0(X_t)\right)\right]+O(\gamma^2), \tag{A.32}$$

$$M_t^{d*,P} = \frac{\beta\xi K_t}{(1+\xi)}\left[1+\gamma\left(\log\left(\frac{\beta^{1+\xi}\xi^{\xi}}{(1+\xi)^{1+\xi}}\right)-\phi_0(X_t)\right)\right]+O(\gamma^2). \tag{A.33}$$

There are a number of important conclusions that can be drawn from the optimal perturbed solutions in Equations (A.32) and (A.33). First, both equations only depend on $\phi_0(X_t)$ and do not depend on $\phi_1(X_t)$, which implies that solving the consumption-investment problem with log-utility is sufficient to fully characterize the optimal perturbed consumption and money holdings up to first order. Secondly, $C_t^{*,P}$ and $M_t^{d*,P}$ are affine functions not only of capital K_t but also of the state vector X_t. This property of the solution will not only render the equilibrium path process of K_t affine, but also implies that optimal inflation dynamics dp_t^*/p_t^* remain affine in the state variables. Next, substituting $C_t^{*,P}$ and $M_t^{d*,P}$ into Equation (9) immediately gives the equilibrium capital process K_t^* in Equation (16). To derive the equilibrium price dynamics in (17), we apply Itô's lemma to the money market clearing condition $M_t^S = p_t^* M_t^{d*}$ and obtain[37]

$$dM_t^S = M_t^{*d} dp_t^* + p_t^* dM_t^{d*} + \mathbb{C}_t\left(dp_t^*, dM_t^{*d}\right). \tag{A.34}$$

[36] Inserting Equation (A.30) into (A.31) gives the optimal money demand in Equation (A.16) from which the optimal consumption in Equation (A.15) can easily be deduced.

[37] Assuming that at $t=0$ markets are cleared. Hence, $p_0^* M_0^{d*} = M_0^S$.

Then using the optimal controls C_t^* and M_t^{d*} and inserting the money market clearing condition from Equation (A.34) yields

$$\frac{dp_t^*}{p_t^*} = \frac{dM_t^S}{M_t^S} - \frac{dK_t^*}{K_t^*} - \mathbb{C}_t\left(\frac{dp_t^*}{p_t^*}, \frac{dK_t^*}{K_t^*}\right). \tag{A.35}$$

Inserting the money supply rule of Equation (10) and the equilibrium capital accumulation process into (A.35) gives the equilibrium price process as in Equation (17). To verify that the guess for the value function $V(\cdot)$ was correct, we substitute the equilibrium values back into the HJB problem in Equation (A.12).

A.3 Proof of Proposition 3

To simplify notation, let $\kappa_t^* = \log(K_t^*) + \beta t$. Using the equilibrium capital accumulation process implies that κ_t^* satisfies

$$d\kappa_t^* = \left(\mu_{K^*}(A_t, v_t) - \frac{1}{2}\sigma_Y^2(k_Y + v_t)\right)dt + \sigma_Y\sqrt{k_Y + v_t}dW_t^Y. \tag{A.36}$$

The Euler condition in Equation (20) can then be expressed as[38]

$$\begin{aligned}
B(t,\tau) &= e^{-\beta\tau}\mathbb{E}_t\left[\frac{U_C(C_{t+\tau}^*, M_{t+\tau}^{d*})}{U_C(C_t^*, M_t^{d*})}\frac{p_t^*}{p_{t+\tau}^*}\right] = e^{-\beta\tau}\mathbb{E}_t\left[\frac{K_t^*}{K_{t+\tau}^*}\frac{p_t^*}{p_{t+\tau}^*}\right] = e^{-\beta\tau}\mathbb{E}_t\left[\frac{\exp\{-\log\left(K_{t+\tau}^*\right)\}}{\exp\{-\log\left(K_t^*\right)\}}\frac{p_t^*}{p_{t+\tau}^*}\right] \\
&= \mathbb{E}_t\left[\frac{\exp\{-\left(\log\left(K_{t+\tau}^*\right) + \beta(t+\tau)\right)\}}{\exp\{-\left(\log\left(K_t^*\right) + \beta t\right)\}}\frac{p_t^*}{p_{t+\tau}^*}\right] = \mathbb{E}_t\left[\frac{\exp\{-\kappa_{t+\tau}^*\}}{\exp\{-\kappa_t^*\}}\frac{p_t^*}{p_{t+\tau}^*}\right].
\end{aligned} \tag{A.37}$$

To solve the problem in Equation (A.37) we follow Ulrich (2013) and Buraschi and Jiltsov (2005) and define

$$f = f(\kappa_t^*, p_t^*, A_t, v_t, m_t, \tau) = \mathbb{E}_t\left[\frac{e^{-\kappa_{t+\tau}^*}}{p_{t+\tau}^*}\right]. \tag{A.38}$$

Conjecturing a log-linear guess for $f(\cdot)$ of the form

$$f(\kappa_t^*, p_t^*, A_t, v_t, m_t, \tau) = \frac{e^{-\kappa_t^*}}{p_t^*}\exp\left\{-b_0(\tau) - b_A(\tau)A_t - b_v(\tau)v_t - b_m(\tau)m_t\right\}. \tag{A.39}$$

If our log-linear guess in Equation (A.39) solves the stochastic problem in (A.38), then it is also the solution to

$$-\frac{\partial f(\cdot,\tau)}{\partial\tau} = \mathcal{A}f(\cdot,\tau), \quad f(\cdot,0) = \frac{e^{-\kappa_t^*}}{p_t^*}. \tag{A.40}$$

[38] From the discussion in Section A.2.2, it suffices to derive the Euler equation for the log-utility case.

The left-hand side of Equation (A.40) is given by

$$\frac{\partial f(\cdot,\tau)}{\partial\tau} = \left[-b_0^{'}(\tau) - b_A^{'}(\tau)A_t - b_v^{'}(\tau)v_t - b_m^{'}(\tau)m_t\right] f(\cdot,\tau). \tag{A.41}$$

Setting $\Theta_t = \{\kappa_t^*, p_t^*, A_t, v_t, m_t\}$, an application of Itô's lemma to the right-hand side of (A.40) gives

$$\mathcal{A}f\left(\kappa_t^*, p_t^*, A_t, v_t, m_t, \tau\right) = \sum_{i\in\Theta}\frac{\partial f}{\partial\Theta^i}\mu(\Theta_t^i)dt + \frac{1}{2}\sum_{i\in\Theta}\frac{\partial^2 f}{\partial\Theta^{i2}}d\langle\Theta^i,\Theta^i\rangle_t + \sum_{\substack{i,j\in\Theta\\ i\neq j}}\frac{\partial^2 f}{\partial\Theta^i\Theta^j}d\langle\Theta^i,\Theta^j\rangle_t. \tag{A.42}$$

Writing out the expression above and substituting the dynamics of $\kappa_t^*,\ p_t^*,\ A_t,\ v_t$, and m_t, we get

$$\begin{aligned}
\mathcal{A}f = {} & \frac{\partial f}{\partial\kappa^*}\left(\mu_{K^*}(A_t,v_t) - \frac{1}{2}\left(\sigma_Y^2(k_Y+v_t)\right)\right) \\
& + \frac{\partial f}{\partial p_t^*}p_t^*\frac{1}{1-\eta_2}\left[\frac{m_t-\eta_1\bar{k}-\eta_2\bar{\pi}}{1-\eta_2} + \frac{\eta_1-1}{1-\eta_2}\left(\mu_{K^*}\left(A_t,v_t\right) - \frac{\sigma_Y^2(k_A+v_t)}{(1-\eta_2)}\right)\right] \\
& + \frac{\partial f}{\partial A}\left(\kappa_A(\theta_A-A_t)+\lambda v_t\right) + \frac{\partial f}{\partial v}\kappa_v(\theta_v-v_t) + \frac{\partial f}{\partial m}\kappa_m(\theta_m-m_t) \\
& + \frac{1}{2}\left[\frac{\partial^2 f}{\partial\kappa^{*2}}\left(\sigma_Y^2(k_Y+v_t)\right) + \frac{\partial^2 f}{\partial p^{*2}}p_t^{*2}\left[\left(\frac{\eta_1-1}{1-\eta_2}\right)^2\left(\sigma_Y^2(k_Y+v_t)\right) + \frac{\sigma_M^2(k_M+v_t)}{(1-\eta_2)^2}\right]\right. \\
& \left. + \frac{\partial^2 f}{\partial A^2}\sigma_A^2(k_A+v_t) + \frac{\partial^2 f}{\partial v^2}\sigma_v^2 v_t + \frac{\partial^2 f}{\partial m^2}\sigma_m^2(k_m+v_t)\right] \\
& + \frac{\partial^2 f}{\partial\kappa^*\partial p^*}p_t^*\frac{(\eta_1-1)}{(1-\eta_2)}\left(\sigma_Y^2(k_Y+v_t)\right) + \frac{\partial^2 f}{\partial p^*\partial m}\frac{(\eta_1-1)}{(1-\eta_2)}\frac{\sigma_M\sigma_m\rho_{Mm}v_t}{(\eta_1-1)} + \frac{\partial^2 f}{\partial p^*\partial v}\frac{(\eta_1-1)}{(1-\eta_2)}\frac{\sigma_M\sigma_v\rho_{Mv}v_t}{(\eta_1-1)} \\
& + \frac{\partial^2 f}{\partial A\partial m}\rho_{Am}k_m\sigma_A\sigma_m + \left(\frac{\partial^2 f}{\partial p^*\partial m}p_t^*\frac{\rho_{Mv}\sigma_v\sigma_m}{(1-\eta_2)} + \frac{\partial^2 f}{\partial A\partial v}\sigma_A\sigma_v\rho_{Av} + \frac{\partial^2 f}{\partial A\partial m}\sigma_A\sigma_m\rho_{Am} + \frac{\partial^2 f}{\partial m\partial v}\sigma_m\sigma_v\rho_{mv}\right)v_t.
\end{aligned}$$

Computing the derivatives and separating variables one obtains the following system of first order asymptotic (Riccati) ODE's:

$$0 = -C_A + \kappa_A b_A(\tau) + b_A^{'}(\tau), \tag{A.43}$$

$$0 = Z_{0v}(\tau) + b_v(\tau)Z_{1v}(\tau) + Z_{2v}b_v^2(\tau) + b_v^{'}(\tau), \tag{A.44}$$

$$0 = -C_m + b_m(\tau)\kappa_m + b_m^{'}(\tau), \tag{A.45}$$

$$0 = -b_0^{'}(\tau) + C_0(\tau), \tag{A.46}$$

subject to $b_A(0) = b_v(0) = b_m(0) = b_0(0) = 0$ and where

$$\begin{aligned}
Z_{0v}(\tau) &= C_v + \left(q_A^2 + 2\gamma\beta\phi_{0A}\right)\left(\frac{\eta_1 - \eta_2}{1 - \eta_2}\right)^2 \frac{(1 - e^{-\kappa_A \tau})^2 \sigma_A^2}{2\kappa_A^2} + \frac{b_m^2(\tau)\sigma_m^2}{2} \\
&\quad + b_A(\tau)\left(b_m(\tau)\sigma_A\sigma_m\rho_{Am} - \lambda\right), \\
Z_{1v}(\tau) &= \kappa_v + \sigma_v\left(b_A(\tau)\sigma_A\rho_{Av} + b_m(\tau)\sigma_m\rho_{mv}\right) + \frac{\sigma_M\sigma_m\rho_{Mm}}{1 - \eta_2}, \\
Z_{2v} &= \sigma_v^2/2,\ H_v(\tau) = \sqrt{4Z_{0v}(\tau)Z_{2v} - Z_{1v}^2(\tau)}, \\
C_0(\tau) &= K_0 + \sum_{i \in \{A,v,m\}} b_i(\tau)\theta_i\kappa_i,
\end{aligned} \tag{A.47}$$

with constants

$$C_A := (q_A + \gamma\beta\phi_{0A})\left(\frac{\eta_1 - \eta_2}{1 - \eta_2}\right), \tag{A.48}$$

$$C_m := \frac{1}{1 - \eta_2}, \tag{A.49}$$

$$C_v := -\frac{\sigma_Y\left(\eta_1^2\sigma_Y - \eta_1(\eta_2\sigma_Y + \sigma_Y - 1) + \eta_2^2\sigma_Y - (\eta_2 - 1)\sigma_Y - 1\right) + \sigma_M^2}{(\eta_2 - 1)^2}, \tag{A.50}$$

where the constant K_0 is a lengthy expression. The time-to-maturity function $C_0(\tau)$ collects all the non-state dependent terms. For the existence of a solution to the bond pricing PDE that excludes arbitrage opportunities, we require that the Riccati equations in (A.44) above satisfy the following periodicity condition

$$4Z_{0v}(\tau)Z_{2v} - Z_{1v}^2(\tau) < 0, \forall\, \tau \geq 0. \tag{A.51}$$

This condition essentially rules out singularities of the solution of the Riccati equation above, i.e., for $\tau \geq 0$, the function $b_v(\tau)$ is continuous in τ.

A.4 Proof of Proposition 4

We derive the first order asymptotic nominal short rate and the market prices of real and nominal risks when the investor has CRRA utility. The nominal short rate is defined as the following limit

$$R_t = \lim_{\tau \to 0} Y(t, \tau) = \lim_{\tau \to 0} -\frac{1}{\tau}\left(\log(B(t, \tau))\right). \tag{A.52}$$

To prove Equation (27) we make repeated use of Bernoulli's rule. For instance, to compute $\lim_{\tau\to 0}\frac{b_0(\tau)}{\tau}$ we set $f(\tau)=\int_0^\tau C_0(u)du$ and $g(\tau)=\tau$. Then, using Leibniz' integral rule we obtain

$$\lim_{\tau\to 0}\frac{f(\tau)}{g(\tau)}=\lim_{\tau\to 0}\frac{f'(\tau)}{g'(\tau)}=\lim_{\tau\to 0}C_0(\tau)=K_0,$$

Similarly we have

$$\lim_{\tau\to 0}\frac{b_A(\tau)}{\tau}=C_A,\quad \lim_{\tau\to 0}\frac{b_v(\tau)}{\tau}=-Z_{0v}(0)=C_v,\quad \lim_{\tau\to 0}\frac{b_m(\tau)}{\tau}=\frac{1}{1-\eta_2}, \tag{A.53}$$

which gives the expression for the short rate in Equation (27). Next, to derive the bond excess return, we apply Itô's rule to the closed-form bond price formula in Equation (21) and obtain the following dynamics

$$\frac{dB(t,\tau)}{B(t,\tau)}=\frac{\partial B(t,\tau)}{\partial t}dt+\sum_{i\in\Theta}\frac{\partial B(t,\tau)}{\partial\Theta^i}d\Theta_t^i+\frac{1}{2}\sum_{i\in\Theta}\frac{\partial^2 B(t,\tau)}{\partial\Theta^{i2}}d\langle\Theta^i,\Theta^i\rangle_t, \tag{A.54}$$

where $\Theta=\{A,v,m\}$. Taking time t conditional expectation on both sides and using Equation (A.44) from above, we obtain after some algebra that the expected infinitesimal bond risk premia is given by

$$RP(t,\tau)=\frac{1}{dt}\mathbb{E}_t\left[\frac{dB(t,\tau)}{B(t,\tau)}-R_tdt\right]=\frac{b_v(\tau)k_m\rho_{Mm}\sigma_m\sigma_M}{\eta_2-1}+\frac{b_v(\tau)(\rho_{Mv}\sigma_M\sigma_v+\rho_{Mm}\sigma_m\sigma_M)}{\eta_2-1}v_t, \tag{A.55}$$

which is the expression in Equation (28). The results for the log-utility agent can be easily derived by simply setting $\gamma=0$ in Proposition (4).

A.5 Model-implied moments

The model-implied unconditional nominal term structure and volatility curve are given by

$$\mathbb{E}[Y(t,\tau)]=\frac{b_0(\tau)}{\tau}+\frac{b_A(\tau)}{\tau}\left(\theta_A+\frac{\lambda\theta_v}{\kappa_A}\right)+\frac{b_v(\tau)}{\tau}\theta_v+\frac{b_m(\tau)}{\tau}\theta_m, \tag{A.56}$$

$$\sqrt{\mathbb{V}[dY(t,\tau)]}=\frac{1}{\tau}\sqrt{\mathbb{V}\left[\frac{dB(t,\tau)}{B(t,\tau)}\right]}=\frac{1}{\tau}\sqrt{\Gamma_0(\tau)+\theta_v\Gamma_1(\tau)}, \tag{A.57}$$

$$\Gamma_0(\tau)=k_A\sigma_A^2b_A^2(\tau)+k_m\sigma_m^2b_m^2(\tau)+k_m\sigma_A\sigma_m\rho_{Am}b_A(\tau)b_m(\tau) \tag{A.58}$$

$$\Gamma_1(\tau)=\sum_{i\in\mathcal{I}}b_i(\tau)^2\sigma_i^2+2\sum_{\substack{i,j\in\mathcal{I}\\ i\neq j}}b_i(\tau)\sigma_ib_j(\tau)\sigma_j\rho_{ij}, \tag{A.59}$$

where $\mathcal{I} = \{A, m, v\}$. To derive the results above, we need the unconditional moments of the factors A_t, m_t, v_t, which we provided in Appendix A.1.

A.6 Model-implied regression coefficient

Bond excess returns over a time period $t, t+h$ and maturity τ are given by

$$r^{E,\tau}_{t,t+h} = \log\left(B(t+h,\tau-h)\right) - \log\left(B(t,\tau)\right) - Y(t,h), \tag{A.60}$$

$$= \log\left(B(t+h,\tau-h)\right) - \log\left(B(t,\tau)\right) + \frac{1}{h}\log\left(B(t,h)\right). \tag{A.61}$$

The theoretical slope coefficient $\beta^{\tau,h}_v$, of a univariate regression of excess returns $r^{E,\tau}_{t,t+h}$ onto policy uncertainty v_t is given by

$$\beta^{\tau,h}_v = \frac{\mathbb{C}\left(r^{E,\tau}_{t,t+h}, v_t\right)}{\mathbb{V}(v_t)}, \tag{A.62}$$

where the denominator is derived in Equation (A.7). We now show the derivation of the covariance term in Equation (A.62). An application of Itô's lemma to Equation (21) shows that the bond price has the dynamics

$$\begin{aligned} dB(t,\tau) &= \frac{\partial B(t,\tau)}{\partial t}dt + \sum_{i\in\Theta}\frac{\partial B(t,\tau)}{\partial \Theta^i}d\Theta^i_t + \frac{1}{2}\sum_{i,j\in\Theta}\frac{\partial^2 B(t,\tau)}{\partial\Theta^i\partial\Theta^j}d\langle\Theta^i,\Theta^j\rangle_t \\ &= B(t,\tau)\left[b^{'}_0(\tau) + \sum_{i\in\Theta}\left(b^{'}_{\Theta^i}(\tau)\Theta^i_t - B(t,\tau)b_{\Theta^i}(\tau)\mu\left(\Theta^i_t\right)\right)\right]dt - B(t,\tau)\sum_{i\in\Theta}b_{\Theta^i}(\tau)\sigma(\Theta^i_t)dW^{\Theta^i}_t \\ &+ \frac{1}{2}B(t,\tau)\left(\sum_{i,j\in\Theta}\rho_{\Theta^i\Theta^j}b_{\Theta^i}(\tau)b_{\Theta^j}(\tau)\sigma(\Theta^i_t)\sigma(\Theta^j_t)\right)dt, \end{aligned} \tag{A.63}$$

where $\mu\left(\Theta^i_t\right)$ is the drift part of the diffusion Θ^i_t where $\Theta_t = \{A_t, v_t, m_t\}$, i.e., $\mu\left(\Theta^1_t\right) = \mu(A_t) = \kappa_A\left(\theta_A - A_t\right) + \lambda v_t$, $\mu\left(\Theta^3_t\right) = \mu(m_t) = \kappa_m\left(\theta_m - m_t\right)$ and $\mu\left(\Theta^2_t\right) = \mu(v_t) = \kappa_v\left(\theta_v - v_t\right)$, respectively. Likewise, $\sigma\left(\Theta^1_t\right) = \sigma(v_t) = \sigma_A(k_A + v_t)$, $\sigma\left(\Theta^2_t\right) = \sigma(v_t) = \sigma_m(k_m + v_t)$ and $\sigma\left(\Theta^3_t\right) = \sigma(v_t) = \sigma_v v_t$. Therefore, using Equation (A.63), we obtain that the log-bond price has the following dynamics

$$\begin{aligned} d\left(\log\left(B(t,\tau)\right)\right) &= \left[b^{'}_0(\tau) + \sum_{i\in\Theta}\left(b^{'}_{\Theta^i}(\tau)\Theta^i_t - b_{\Theta^i}(\tau)\mu\left(\Theta^i_t\right)\right)\right]dt - \sum_{i\in\Theta}b_{\Theta^i}(\tau)\sigma(\Theta^i_t)dW^{\Theta^i}_t \\ &= \mu_B(t,\tau)dt - \sum_{i\in\Theta}b_{\Theta^i}(\tau)\sigma(\Theta^i_t)dW^{\Theta^i}_t. \end{aligned} \tag{A.64}$$

Using Equation (A.64), we can rewrite

$$\log\left(B(t+h,\tau-h)\right)-\log\left(B(t,\tau)\right)=\int_t^{t+h} d\left(\log\left(B(s,t+\tau-s)\right)\right)$$

$$-\frac{1}{h}\log\left(B(t,h)\right)=\frac{1}{h}\left(\log\left(B(t+h,0)\right)-\log\left(B(t,h)\right)\right)=\frac{1}{h}\int_t^{t+h} d\left(\log(B(s,t+h-s))\right),$$

to express the excess return in Equation (A.61) as

$$\begin{aligned}r^{E,\tau}_{t,t+h}=&\int_t^{t+h}\left[\mu_B(s,t+\tau-s)-\frac{1}{h}\mu_B(s,t+h-s)\right]ds\\&-\sum_{i\in\Theta}\int_t^{t+h}\left[b_{\Theta^i}(t+\tau-s)-\frac{1}{h}b_{\Theta^i}(t+h-s)\right]\sigma(\Theta^i_t)dW^{\Theta^i}_s.\end{aligned}\tag{A.65}$$

Since the Brownian motions $W^{\Theta^i}_t$ are independent for $s>t$, we can focus on the drift part in (A.65). Then, the covariance between excess bond returns over the time interval t to $t+h$ and policy uncertainty v_t is given by

$$\begin{aligned}\mathbb{C}\left(r^{E,\tau}_{t,t+h},v_t\right)=&\,\mathbb{C}\Bigg(\int_t^{t+h}\sum_{i\in\Theta}\Bigg[\left(b'_{\Theta^i}(t+\tau-s)-\frac{b'_{\Theta^i}(t+h-s)}{h}\right)\Theta^i_s\\&-\left(b_{\Theta^i}(t+\tau-s)-\frac{b_{\Theta^i}(t+h-s)}{h}\right)\mu\left(\Theta^i_s\right)\Bigg]ds,v_t\Bigg)\\=&\int_t^{t+h}\sum_{i\in\Theta}\left(b'_{\Theta^i}(t+\tau-s)-\frac{b'_{\Theta^i}(t+h-s)}{h}\right)\mathbb{C}\left(\Theta^i_s,v_t\right)ds\\&-\int_t^{t+h}\sum_{i\in\Theta}\left(b_{\Theta^i}(t+\tau-s)-\frac{b_{\Theta^i}(t+h-s)}{h}\right)\mathbb{C}\left(\mu\left(\Theta^i_s\right),v_t\right)ds,\end{aligned}\tag{A.66}$$

where the second equality follows due to Fubini's theorem. Equation (A.66) implies that in order to compute the unconditional covariance between excess returns and economic policy uncertainty v_t, we need the covariances between the state variables Θ_s and v_t. These are provided in Section (A.1).

References

Ang, A., and M. Piazzesi, 2003, "A No-Arbitrage Vector Autoregression of Term Structure Dynamics with Macroeconomic and Latent Variables," *Journal of Monetary Economics*, 50, 745–787.

Baker, S., N. Bloom, and S. J. Davis, 2016, "Measuring Economic Policy Uncertainty," *Quarterly Journal of Economics*, 131(4), 1593–1636.

Baker, S. R., and N. Bloom, 2013, "Does Uncertainty Reduce Growth? Using Disasters as Natural Experiments," Working Paper 19475, National Bureau of Economic Research.

Balduzzi, P., E. Elton, and T. Green, 2001, "Economic News and Bond Prices: Evidence from the U.S. Treasury Market.," *Journal of Financial and Quantitative Analysis*, 36, 523–543.

Bansal, R., and I. Shaliastovich, 2013, "A Long-Run Risks Explanation of Predictability Puzzles in Bond and Currency Markets," *Review of Financial Studies*, 26(1), 1–33.

Bansal, R., and A. Yaron, 2004, "Risks for the Long Run: A Potential Resolution of Asset Pricing Puzzles," *Journal of Finance*, 59, 1481–1509.

Bekaert, G., E. Engstrom, and Y. Xing, 2009, "Risk, Uncertainty, and Asset Prices," *Journal of Financial Economics*, 91(1), 59–82.

Bekaert, G., C. R. Harvey, C. T. Lundblad, and S. Siegel, 2014, "Political Risk Spreads," *Journal of International Business Studies*, 45(4), 471–493.

Belo, F., V. Gala, and J. Li, 2013, "Government Spending, Political Cycles and the Cross Section of Stock Returns," *Journal of Financial Economics*, 107, 305–324.

Bialkowski, J., K. Gottschalk, and T. Wisnieski, 2008, "Stock Market Volatility around National Elections," *Journal of Banking and Finance*, 32, 1941–1953.

Bloom, N., 2009, "The Impact of Uncertainty Shocks," *Econometrica*, 77(3), 623–685.

Bond, P., and I. Goldstein, 2015, "Government Intervention and Information Aggregation by Prices," *Journal of Finance*, 70(6), 2777–2812.

Boutchkova, M. K., H. Doshi, A. Durnev, and A. Molchanov, 2012, "Precarious Politics and Return Volatility," *Review of Financial Studies*, 25, 1111–1154.

Brogaard, J., and A. Detzel, 2015, "The Asset-Pricing Implications of Government Economic Policy Uncertainty," *Management Science*, 61(1), 3–18.

Buraschi, A., and A. Jiltsov, 2005, "Inflation Risk Premia and the Expectation Hypothesis," *Journal of Financial Economics*, 75, 429–490.

——— , 2007, "Habit Formation and Macroeconomic Models of the Term Structure of Interest Rates," *Journal of Finance*, 6, 3009–30063.

Campbell, J. Y., and R. J. Shiller, 1991, "Yield Spreads and Interest Rate Movements: A Bird's Eye View," *Review of Economic Studies*, 58, 495–514.

Cochrane, J., and M. Piazzesi, 2005, "Bond Risk Premia," *American Economic Review*, 95, 138–160.

Croce, M., H. Kung, T. Nguyen, and L. Schmid, 2012, "Fiscal Policies and Asset Prices," *Review of Financial Studies*, 25, 2636–2672.

Croce, M., T. Nguyen, and L. Schmid, 2012, "The Market Price of Fiscal Uncertainty," *Journal of Monetary Economics*, 59, 401–4016.

David, A., and P. Veronesi, 2014, "Investors' and Central Bank's Uncertainty Embedded in Index Options," *Review of Financial Studies*, 27(6), 1661–1716.

de Goeij, P., and W. Marquering, 2006, "Macroeconomic Announcements and Asymmetric Volatility in Bond Returns," *Journal of Banking and Finance*, 30(10), 2659–2680.

Duffie, D., and R. Kan, 1996, "A Yield Factor Model of Interest Rates," *Mathematical Finance*, 6, 379–406.

Durnev, A., 2010, "The Real Effects of Political Uncertainty: Elections and Investment Sensitivity to Stock Prices," *Working Paper*, University of Iowa.

Evans, C., and D. Marshall, 2007, "Economic determinants of the nominal treasury yield curve," *Journal of Monetary Economics*, 54, 1986–2003.

Fama, E., and K. French, 1989, "Business Conditions and Expected Returns on Stocks and Bonds," *Journal of Financial Economics*, 25, 23–49.

Fisher, S., and F. Modigliani, 1978, "Towards an Understanding of the Real Effects and Costs of Inflation," *Weltwirtschaftliches Archiv*, 114, 810–832.

Fleming, M., and M. Piazzesi, 2005, "Monetary Policy Tick-by-Tick," *Working Paper*.

Gallmeyer, M. F., B. Hollifield, F. Palomino, and S. E. Zin, 2007, "Arbitrage-Free Bond Pricing with Dynamic Macroeconomic Models," Working Paper 13245, National Bureau of Economic Research.

Gulen, H., and M. Ion, 2012, "Policy Uncertainty and Corporate Investment," *Working Paper*.

Gürkaynak, R. S., B. Sack, and E. Swanson, 2005a, "Do Actions Speak Louder Than Words? The Response of Asset Prices to Monetary Policy Actions and Statements," *International Journal of Central Banking*, pp. 55–93.

——— , 2005b, "The Sensitivity of Long-Term Interest Rates to Economic News: Evidence and Implications for Macroeconomic Models," *The American Economic Review*, 95, 425–426.

Hansen, L. P., J. C. Heaton, and N. Li, 2008, "Consumption Strikes Back? Measuring Long-Run Risk," *Journal of Political Economy*, 116(2), pp. 260–302.

Huang, T., F. Wu, J. Yu, and B. Zhang, 2015, "International political risk and government bond pricing," *Journal of Banking and Finance*, 55, 393–405.

Jeanblanc, M., M. Yor, and M. Chesney, 2009, *Mathematical Methods for Financial Markets*. Springer.

Joslin, S., M. Priebisch, and K. Singleton, 2014, "Risk Premiums in Dynamic Term Structure Models with Unspanned Macro Risks," *Journal of Finance*, 69, 1197–1233.

Julio, J., and Y. Yook, 2012, "Political Uncertainty and Corporate Investment Cycles," *Journal of Finance*, 67, 45–83.

Kelly, B., L. Pástor, and P. Veronesi, 2016, "The Price of Political Uncertainty: Theory and Evidence from the Option Market," *Journal of Finance*, 71(5), 2417–2480.

Kogan, L., and R. Uppal, 2001, "Risk Aversion and Optimal Portfolio Policies in Partial and General Equilibrium Economies," Working paper, Sloan School of Management, MIT.

Kuttner, K. N., 2001, "Monetary Policy Surprises and Interest Rates: Evidence from the Fed Funds Futures Market," *Journal of Monetary Economics*, 47, 523–544.

Litterman, R. B., and J. Scheinkman, 1991, "Common Factors Affecting Bond Returns," *Journal of Fixed Income*, 1, 54–61.

Ludvigson, S., and S. Ng, 2009, "Macro Factors in Bond Risk Premia," *Review of Financial Studies*, 22, 5027–5067.

Malkhozov, A., P. Mueller, A. Vedolin, and G. Venter, 2016, "Mortgage risk and the yield curve," *Review of Financial Studies*, 29, 1220–1253.

Newey, W., and K. West, 1994, "Automatic Lag Selection in Covariance Estimation," *Review of Economic Studies*, 61, 631–653.

Øksendal, B., 2003, *Stochastic Differential Equations*. Springer, Berlin, Heidelberg.

Pastor, L., and P. Veronesi, 2012, "Uncertainty about Government Policy and Stock Prices," *Journal of Finance*, 4, 1219–1264.

——— , 2013, "Political Uncertainty and Risk Premia," *Journal of Financial Economics*, 110, 520–545.

Pennacchi, G., 1991, "Identifying the Dynamics of Real Interest Rates and Inflation Evidence Using Survey Data," *Review of Financial Studies*, 4, 53–86.

Piazzesi, M., 2005, "Bond Yields and the Federal Reserve," *Journal of Political Economy*, 112, 311–344.

Piazzesi, M., and M. Schneider, 2006, "Equilibrium Yield Curves," Working Paper 12609, National Bureau of Economic Research.

Pindyck, R. S., and A. Solimano, 1993, "Economic Instability and Aggregate Investment," Working Paper 4380, National Bureau of Economic Research.

Rodrik, D., 1991, "Policy Uncertainty and Private Investment in Developing Countries," *Journal of Development Economics*, 36, 229–242.

Scotti, C., 2016, "Surprise and uncertainty indexes: Real-time aggregation of real-activity macro-surprises," *Journal of Monetary Economics*, 82, 1 – 19.

Shiller, R., 1979, "The Volatility of Long-Term Interest Rates and Expectations Models of the Term Structure," *Journal of Political Economy*, 87, 1190–1219.

Ulrich, M., 2013, "Inflation Ambiguity and the Term Structure of U.S. Government Bonds," *Journal of Monetary Economics*, 60, 295–309.

Wachter, J. A., 2006, "A Consumption-based Model of the Term Structure of Interest Rates," *Journal of Financial Economics*, 79(2), 365–399.

Wright, J., 2012, "What does Monetary Policy do to Long-term Interest Rates at the Zero Lower Bound?," *The Economic Journal*, 122, 447–66.

Xiong, W., and H. Yan, 2010, "Heterogenous Expectations and Bond Markets," *Review of Financial Studies*, 23, 1433–1466.

Tables

			Panel A: Economic policy uncertainty v_t				
κ_v	0.169	θ_v	1.072	σ_v	0.177		
	(0.041)		(0.069)		(0.014)		
			Panel B: Matched and fixed parameters				
		$\bar{k}$	0.156	$\bar{\pi}$	0.165		
			Panel C: Calibrated parameters				
β	0.026	k_A	0.375	ρ_{Am}	-0.884	κ_A	3.358
ξ	0.463	k_M	4.311	ρ_{Av}	0.168	θ_A	9.791
γ	-0.466	k_m	0.119	ρ_{Mm}	0.399	σ_A	2.354
δ	0.075	k_Y	4.903	ρ_{Mv}	-0.240	κ_m	0.363
η_1	-1.012	μ_Y	0.497	ρ_{mv}	-0.566	θ_m	10.091
η_2	-2.242	σ_Y	0.638	λ	-5.275	σ_m	8.252
q_A	2.959	σ_M	2.548				

Table 1: Summary of parameter values. In Panel A, we report the maximum likelihood estimates for the policy uncertainty process v_t. We proxy v_t using the EPU index, which we scale by dividing it by 100. In brackets, we report the corresponding asymptotic robust standard errors ('Sandwich estimator') based on the outer product of the Jacobian of the log-likelihood function. In Panel B, we report the matched and fixed parameters. The remaining parameters in Panel C are calibrated to match simultaneously the unconditional yield and bond volatility curves. Estimation period is January 1990 to September 2015 (309 data points) using monthly data.

	1Y	2Y	3Y	5Y	7Y	10Y
$\frac{1}{\tau}b_v(\tau)$	-0.008	-0.015	-0.024	-0.041	-0.054	-0.065

Table 2: The table reports the calibrated values for the policy uncertainty factor loadings $b_v(\tau)/\tau$, obtained using the parameters as given in Table 1. The function $b_v(\tau)$ solves the Ricatti Equation in (23).

Mat	EPU	Economic conditions VIX	CFN	TIV	Financial variables DY	TS	Macroeconomic controls IP	CPI	R^2(adj)
1Y	−3.715***	–	–	–	–	–	–	–	0.28
	−4.047***	0.020	0.038	1.201	–	–	–	–	0.30
	−2.727***	0.001	0.680***	2.371**	−1.133***	1.916***	–	–	0.70
	−2.843***	0.023	0.922***	2.112**	−1.084***	1.955***	−0.792	1.885**	0.71
2Y	−3.693***	–	–	–	–	–	–	–	0.27
	−3.912***	0.004	0.095	1.843*	–	–	–	–	0.30
	−3.000***	0.003	0.714***	2.508**	−0.967***	1.994***	–	–	0.66
	−3.127***	0.028	0.970***	2.220**	−0.912***	2.035***	−0.839	2.095**	0.67
3Y	−3.529***	–	–	–	–	–	–	–	0.27
	−3.676***	−0.008	0.108	2.237**	–	–	–	–	0.31
	−3.070***	0.004	0.693***	2.534**	−0.823***	1.987***	–	–	0.63
	−3.202***	0.028	0.931***	2.237**	−0.767***	2.024***	−0.789	2.181***	0.64
5Y	−3.045***	–	–	–	–	–	–	–	0.23
	−3.060***	−0.028	0.130	2.770***	–	–	–	–	0.29
	−2.919***	0.002	0.650***	2.521**	−0.592***	1.928***	–	–	0.59
	−3.052***	0.025	0.844**	2.222**	−0.537***	1.957***	−0.657	2.222***	0.60
7Y	−2.632***	–	–	–	–	–	–	–	0.20
	−2.595***	−0.033	0.153	2.930***	–	–	–	–	0.27
	−2.681***	0.004	0.628***	2.427***	−0.465***	1.845***	–	–	0.55
	−2.818***	0.027	0.792**	2.119**	−0.409***	1.868***	−0.569	2.314***	0.57
10Y	−2.116***	–	–	–	–	–	–	–	0.15
	−1.979***	−0.047**	0.141	3.098***	–	–	–	–	0.25
	−2.284***	−0.002	0.578***	2.345***	−0.349***	1.790***	–	–	0.55
	−2.412***	0.019	0.722**	2.057**	−0.296**	1.809***	−0.504	2.169***	0.56

Table 3: Yield curve regressions. The table displays slope coefficients of the regression of yields on EPU_t (EPU) and different controls for economic conditions (VIX, CFN, TIV), financial variables (DY, TS), and macroeconomic controls (IP, CPI). The yield maturities are 1, 2, 3, 5, 7, and 10 years. The last column reports the adjusted R^2-values. By ***, **, * we denote 1%, 5%, and 10% statistical significance. By ***, **, * we denote 1%, 5%, and 10% statistical significance according to the HAC-robust t-statistics. The definitions of the control variables are given in Section 3.2. The sample period is January 1990 to September 2015.

		Economic conditions			Financial variables		Macroeconomic controls		
Mat	EPU	VIX	CFN	TIV	DY	TS	IP	CPI	R^2(adj)
1Y	1.973***	–	–	–	–	–	–	–	0.33
	1.954***	0.010	−0.010	−0.582	–	–	–	–	0.34
	1.549***	0.018	−0.169	−0.972	0.308***	−0.444***	–	–	0.45
	1.618***	0.009	−0.175	−0.817	0.281***	−0.441***	0.056	−1.213**	0.46
2Y	2.013***	–	–	–	–	–	–	–	0.43
	1.941***	0.001	−0.080	−0.093	–	–	–	–	0.43
	1.707***	0.001	−0.243***	−0.259	0.252***	−0.528***	–	–	0.55
	1.753***	−0.006	−0.256*	−0.157	0.234***	−0.527***	0.066	−0.793*	0.56
3Y	1.721***	–	–	–	–	–	–	–	0.45
	1.618***	0.003	−0.067	0.077	–	–	–	–	0.45
	1.503***	−0.001	−0.197**	0.036	0.176***	−0.448***	–	–	0.56
	1.543***	−0.005	−0.152	0.126	0.161***	−0.437***	−0.111	−0.733**	0.57
5Y	1.271***	–	–	–	–	–	–	–	0.45
	1.140***	0.007	−0.068	0.007	–	–	–	–	0.47
	1.100***	0.003	−0.147**	0.021	0.097**	−0.285***	–	–	0.54
	1.131***	−0.001	−0.127	0.090	0.085**	−0.279***	−0.039	−0.559**	0.54
7Y	0.953***	–	–	–	–	–	–	–	0.42
	0.825***	0.007	−0.054	0.018	–	–	–	–	0.44
	0.816***	0.004	−0.110**	0.053	0.062**	−0.211**	–	–	0.50
	0.845***	0.001	−0.090	0.119	0.051*	−0.205**	−0.044	−0.529***	0.51
10Y	0.673***	–	–	–	–	–	–	–	0.38
	0.538***	0.010**	−0.037	0.029	–	–	–	–	0.43
	0.531***	0.007	−0.077*	0.052	0.044**	−0.148**	–	–	0.48
	0.554***	0.005	−0.051	0.105	0.035*	−0.141**	−0.064	−0.428***	0.49

Table 4: Bond yield volatility regressions. The table displays slope coefficients of the regression of yield volatilities on EPU_t (EPU) and different controls for economic conditions (VIX, CFN, TIV), financial variables (DY, TS), and macroeconomic controls (IP, CPI). The yield maturities are 1, 2, 3, 5, 7, and 10 years. The last column reports the adjusted R^2-values. By ***, **, * we denote 1%, 5%, and 10% statistical significance. By ***, **, * we denote 1%, 5%, and 10% statistical significance according to the HAC-robust t-statistics. The definitions of the control variables are given in Section 3.2. The sample period is January 1990 to September 2015.

		Term structure controls				Economic conditions			Financial variables		Macroeconomic controls		
Mat	EPU	PC1	PC2	PC3	CP	VIX	CFN	TIV	DY	TS	IP	CPI	R^2(adj)
2Y	0.181	–	–	–	–	–	–	–	–	–	–	–	0.01
	−0.415***	−2.337	−1.324	−0.031	0.224***	–	–	–	–	–	–	–	0.29
	−0.377**	−2.972	−1.335	−0.020	0.203***	−0.006	0.029	0.683***	–	–	–	–	0.34
	−0.021	−3.486***	−0.697	0.778*	−0.356***	0.001	0.001	0.159	0.644***	0.190***	–	–	0.66
	−0.089	−4.096***	−1.203*	1.039**	−0.269***	−0.006	−0.090	0.348**	0.555***	0.175**	0.113	0.301	0.70
3Y	0.794***	–	–	–	–	–	–	–	–	–	–	–	0.08
	−0.118	0.322	1.254	3.111***	0.417***	–	–	–	–	–	–	–	0.47
	−0.167	−0.373	1.460	2.968***	0.407***	0.002	0.032	0.547*	–	–	–	–	0.49
	0.237	−0.879	2.203**	3.806***	−0.211**	0.010	−0.002	−0.019	0.714***	0.191**	–	–	0.62
	0.316**	−1.551	1.643*	4.210***	−0.260***	−0.006	−0.187	0.272	0.781***	0.124	0.247	0.259	0.72
5Y	3.106***	–	–	–	–	–	–	–	–	–	–	–	0.21
	0.503*	−9.050***	3.983**	9.170***	1.192***	–	–	–	–	–	–	–	0.73
	0.267	−9.995***	4.656**	8.689***	1.204***	0.020	0.044	0.367	–	–	–	–	0.73
	0.691**	−10.703***	5.390***	9.723***	0.521***	0.030	0.014	−0.292	0.787***	0.259*	–	–	0.76
	1.152***	−11.904***	5.151***	10.552***	0.150	−0.002	−0.365	0.253	1.214***	0.096	0.776**	−0.141	0.82
7Y	4.212***	–	–	–	–	–	–	–	–	–	–	–	0.30
	1.568***	−4.420	−3.691*	9.251***	1.125***	–	–	–	–	–	–	–	0.64
	1.071***	−4.579	−2.201	8.322***	1.196***	0.055**	0.066	−0.853	–	–	–	–	0.64
	1.312***	−4.576	−1.681	8.553***	0.887***	0.056**	0.035	−1.078*	0.359*	0.012	–	–	0.64
	2.244***	−5.882	−1.484	9.547***	0.066	0.005	−0.559*	−0.283	1.281***	−0.271	1.323**	−0.742	0.71
10Y	7.040***	–	–	–	–	–	–	–	–	–	–	–	0.35
	3.669***	−9.431	14.993***	1.083	1.537***	–	–	–	–	–	–	–	0.66
	2.747***	−8.960	16.897***	−0.400	1.690***	0.091**	−0.057	−2.561**	–	–	–	–	0.67
	2.265***	−8.529	15.967***	−1.247	2.394***	0.084**	−0.010	−1.950**	−0.814***	−0.170	–	–	0.67
	4.038***	−9.223	17.884***	−0.059	0.739**	0.021	−0.782	−1.257	1.036***	−0.595*	2.013*	−2.147	0.69

Table 5: Bond risk premium regressions. The table displays slope coefficients of regressing the bond excess return r_{t+1}^{E,τ_i} on EPU_t (EPU) and on a set of different controls: the first three principal components PC1, PC2, and PC3, the Cochrane-Piazzesi factor (CP), economic conditions (VIX, CFN, TIV), financial variables (DY, TS), and macroeconomic controls (IP, CPI). The yield maturities are 1, 2, 3, 5, 7, and 10 years. The last column reports the adjusted R^2-values. By ***, **, * we denote 1%, 5%, and 10% statistical significance. By ***, **, * we denote 1%, 5%, and 10% statistical significance according to the HAC-robust t-statistics. The definitions of the control variables are given in Section 3.2. The sample period is January 1990 to September 2015.

Figures

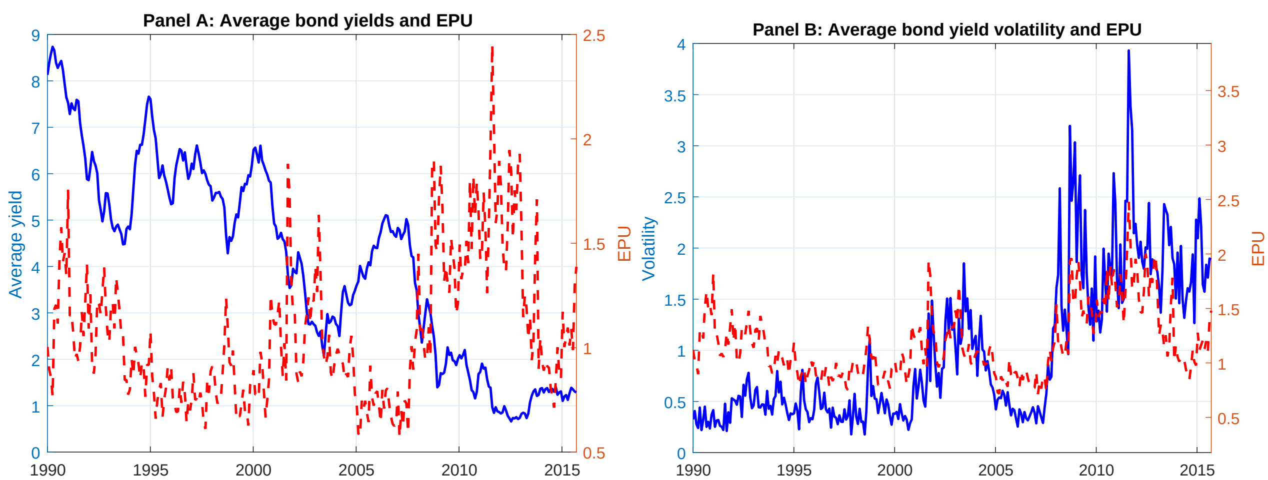

Figure 1: Average US Treasury bond yields (Panel A, solid line) and realized yield volatility (Panel B, solid line) with maturity $\tau = 1$Y, 2Y, 3Y, 5Y, 7Y and 10Y and the economic policy uncertainty index (EPU, dashed line) as constructed by Baker, Bloom, and Davis (2016). The sample period ranges from January 1990 until September 2015.

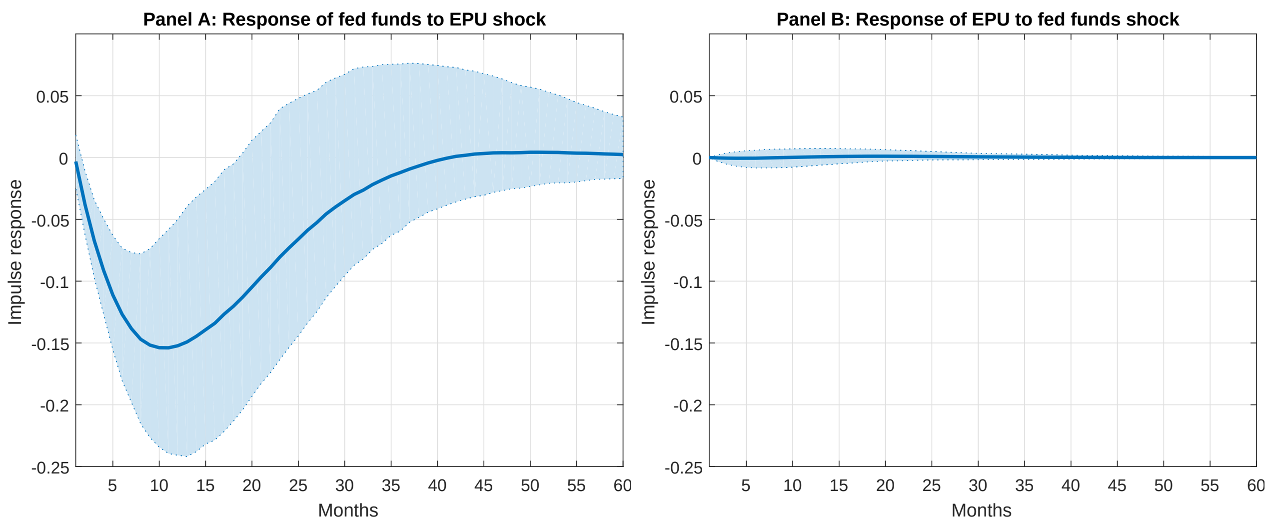

Figure 2: The figure plots the impulse response functions of a shock to the EPU on the short rate (Panel A) and a shock to the short rate on the EPU (Panel B). The short rate is approximated by the three-month T-bill rate. The impulse response functions are based on a bivariate VAR model including the EPU and the effective funds rate. The data sample spans the period from January 1990 to September 2015. The shaded are corresponds to the 95% confidence interval.

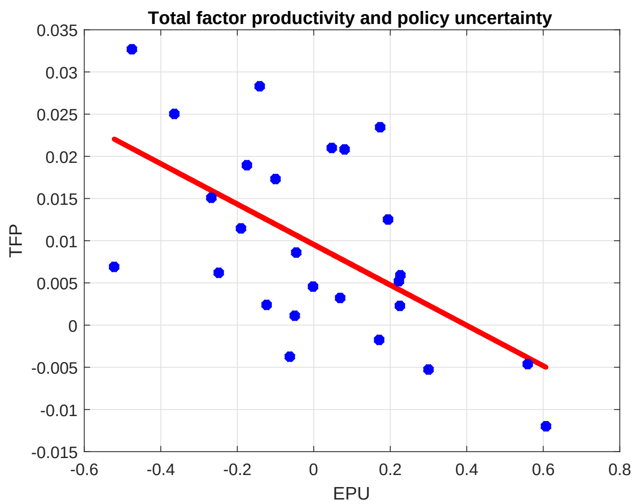

Figure 3: Total factor productivity and the EPU index. The solid line represents the univariate regression line of TFP onto EPU. The estimated slope coefficient is $\hat{\beta}_{EPU} = -0.024$ with corresponding HAC-robust standard error of $\mathrm{SE}\left(\hat{\beta}_{EPU}\right) = 0.0068$. The data are annually and range from 1990 until 2015. Source: Bureau of Labor Statistics and time series name is Multi-factor productivity.

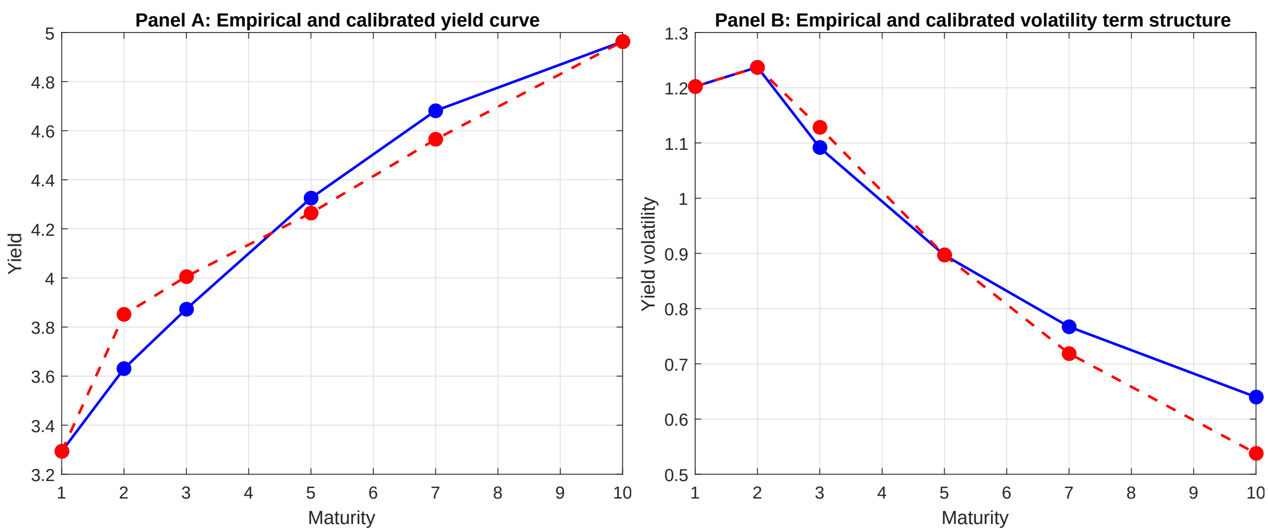

Figure 4: Empirical and fitted affine yield curve model. In Panel A we plot the empirical unconditional nominal yield curve based on monthly zero-coupon bonds (solid line) with the model-implied yield curve (dashed line). Panel B compares the model-implied bond volatility term-structure (dashed line) to the empirical unconditional realized volatility term structure (solid line). Unconditional realized volatility is computed using monthly log-yield changes. The model-implied yields and volatilities are based on the parameters in Table 1. Sample period ranges from January 1990 until September 2015.

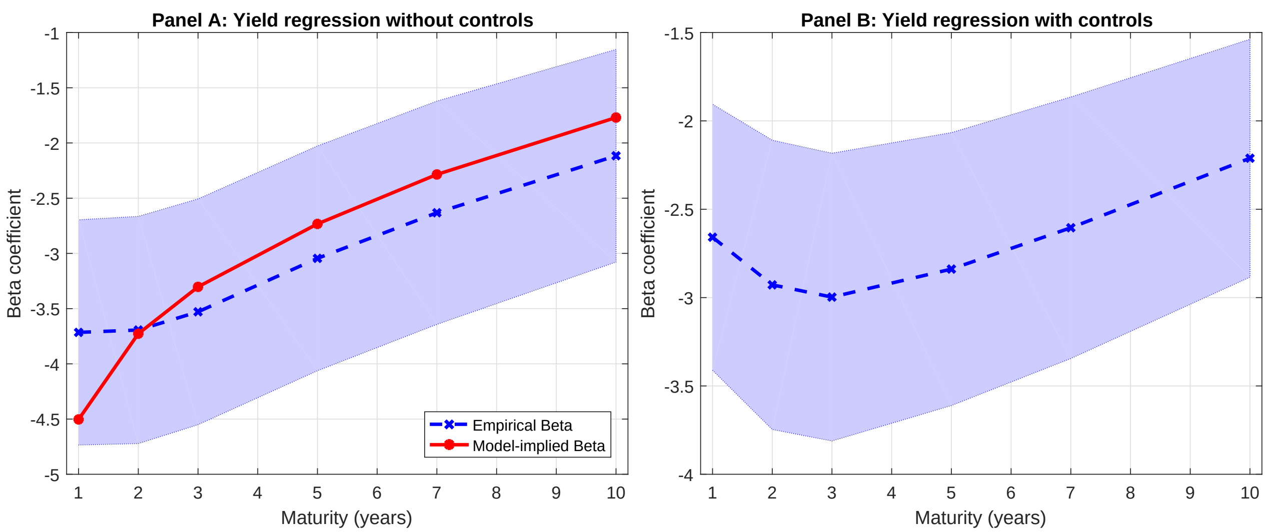

Figure 5: Yield curve regressions. Panel A compares the empirical slope coefficient of the univariate regression of yields $Y(t, \tau)$ onto the EPU (dashed line) with the model-implied regression coefficient (solid line) given in Equation (37). The yield maturities are 1, 2, 3, 5, 7, and 10 years. Panel B plots the regression coefficient we obtain by including all the controls for economic, financial, and macroeconomic conditions (dashed line). Shaded areas represent HAC-robust 95% confidence bounds. The sample period ranges from January 1990 until September 2015.

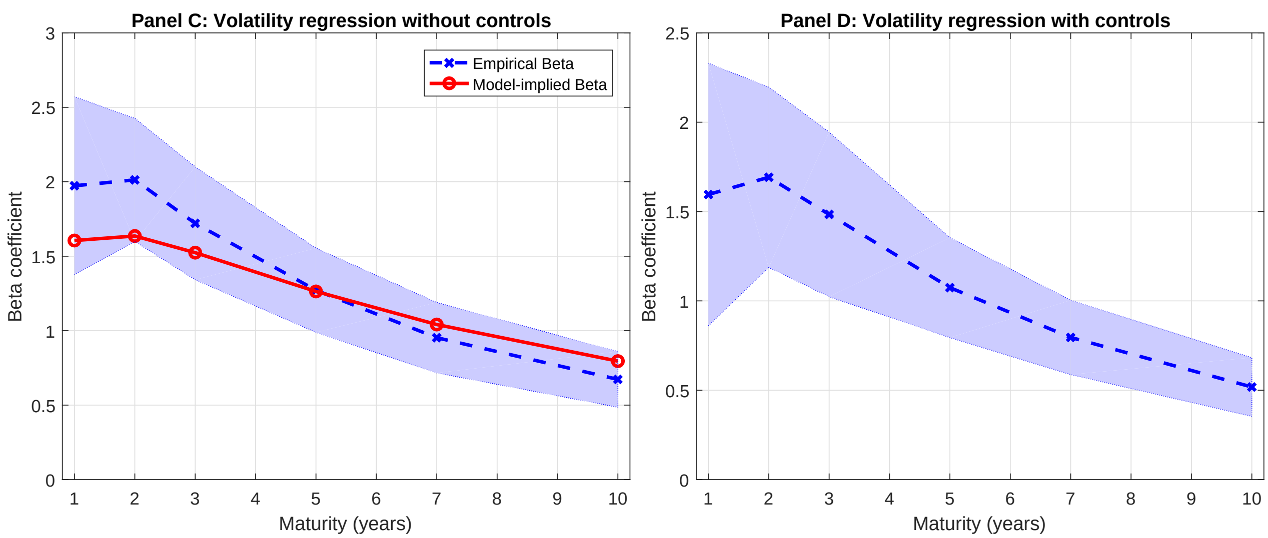

Figure 6: Bond yield volatility regressions. Panel A compares the empirical slope coefficient of the univariate regression of conditional bond yield volatility $\mathcal{V}_t(Y(t,\tau))$ onto the EPU (dashed line) with the model-implied regression coefficient (solid line) defined in Equation (38). The yield maturities are 1, 2, 3, 5, 7, and 10 years. Panel B plots the regression coefficient we obtain by including all the controls for economic, financial, and macroeconomic conditions (dashed line). Shaded areas represent HAC-robust 95% confidence bounds. The sample period ranges from January 1990 until September 2015.

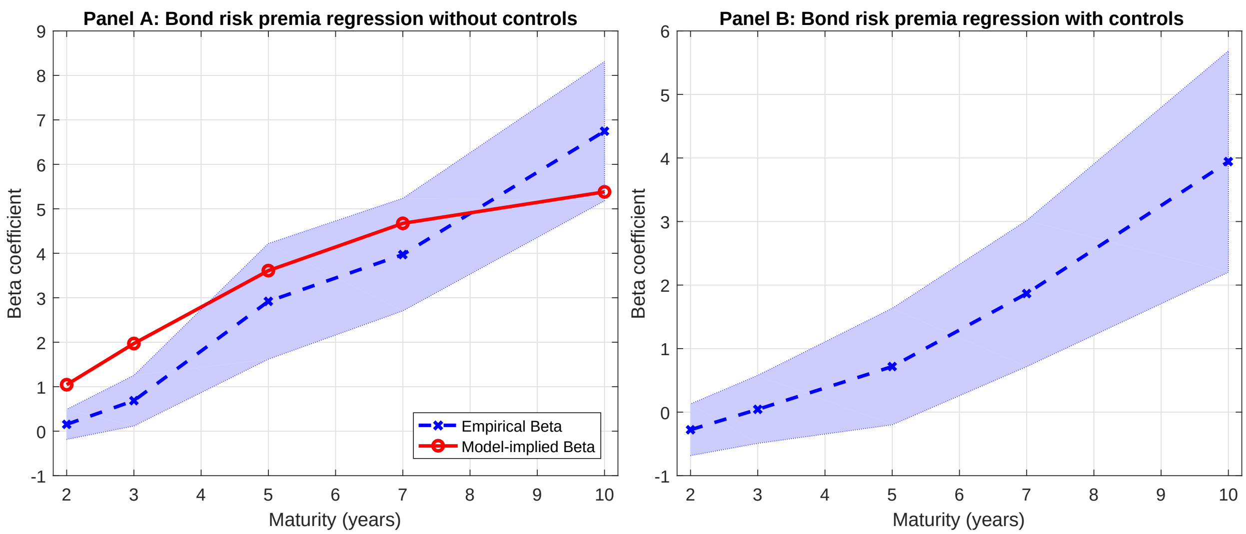

Figure 7: Bond risk premium regressions. The figure compares the empirical slope coefficient of the univariate regression of bond excess returns as defined in (43) on the EPU (dashed line) with the model-implied regression coefficient (solid line) given in Equation (39). The yield maturities are 1, 2, 3, 5, 7, and 10 years. Panel B plots the regression coefficient we obtain by including all the controls for economic, financial, and macroeconomic conditions (dashed line). Shaded areas represent HAC-robust 95% confidence bounds. The sample period ranges from January 1990 until September 2015.